External Threat Causes Nationalist Behavior[*]

Benjamin Helms[†] Sonal Pandya[‡] Rajkumar Venkatesan[§]

February 14, 2020

Abstract

A growing body of experimental findings point to identity as a key driver of political attitudes. External validation of this evidence, necessary for coherent theories grounded in behavioral micro foundations, is lacking owing the absence of appropriate research designs. We demonstrate that external threats have a causal effect on national identification using a novel behavioral measure of change in national identification: weekly sales of American-sounding supermarket brands. During 2003-2006, the market share of American-sounding brands increased in supermarkets following the death of a local solider during the Iraq War. Cumulative exposure to casualties further increases the market share of American-sounding brands over time, indicating that casualties can have enduring effects on national identification. Stores with high proportions of native-born, less educated customers saw sharper increases; stores with greater racial diversity saw relatively smaller increases. These findings establish, with a high degree of external validity, drivers of change in national identification and provides insights into the current rise of nationalist politics.

[*]We received helpful feedback from Luca Cian and seminar participants at Colorado, Georgetown, Stanford, Texas A&M, Texas Tech, UC Davis, and UC San Diego. Doug Kriner, Francis Shen, and Mike Koch generously shared data. We gratefully acknowledge financial support from the Bankard Fund for Political Economy. Pandya also thanks the Hoover Institution National Fellows Program. Adam Hughes provided excellent research assistance.

[†]PhD Candidate, Department of Politics, University of Virginia. benhelms@virginia.edu.

[‡]**Corresponding Author** Associate Professor, Department of Politics, University of Virginia. +1-434-243-1573. spandya@virginia.edu.

[§]Professor, Darden School of Business, University of Virginia. venkatesanr@darden.virginia.edu.

1 Introduction

The behavioral revolution in Political Science, marked by a renewed focus on individual-level processes and an embrace of experimental methods, has upended long-established theoretical microfoundations. Among its many contributions, this research reveals much richer origins of political attitudes, casting aside assumptions about material motives with compelling evidence of non-material motives like identity.

External validity, whether experimental insights manifest in mass behavior, remains the fundamental challenge to parlaying experimental insights into more coherent theory. The challenge persists owing to the lack of empirical strategies to establish the casual effects of non-material (i.e. difficult to observe) motives for mass behavior. For example, several studies consider how economic hardship (Colantone and Stanig, 2018), war (Baum, 2002), and cultural pluralism (Inglehart and Norris, 2016) fuel support for nationalist politicians. No doubt experimental findings help motivate these studies, but extant research designs cannot directly test whether the hypothesized drivers cause nationalism or whether that nationalism has a causal effect on vote choice. Outside of an experimental setting, a myriad of other factors might produce outcomes inconsistent with experimental findings. Nor are experiments well-suited to evaluate sources of change in non-material motives or the endogeneity of such motives. We are left with a large gap between individual-level findings on expressed/revealed attitudes and infrequent, albeit important, aggregate behavioral outcomes like party vote shares whose underlying drivers cannot be directly evaluated. Bridging this gap would not only produce more coherent theory, but would also furnish sharper insights into the origins of contemporary nationalism.

We provide a novel empirical approach to estimating the causal drivers and consequences of American national identity: weekly change in the market share of American-sounding supermarket brands. Nationality-based branding links products to consumers' associations with the implied nation. Consumers purchase brands consistent with their most important social identities (Escalas and Bettman, 2005; Muniz and O'Guinn, 2001), and that bolster their self-esteem (Shachar et al., 2011). Supermarket purchasing is a frequent, consistent, and nearly universal behavior in the United States (Kahn and Schmittlein, 1989). These purchases are more likely to reflect cognitive and emotional processes that operate outside conscious awareness because consumers rely more heavily on heuristics when purchasing low-cost, nondurable goods (Maheswaran, 1994). Purchases are unobtrusively measured, span thousands of brands, and are available for nationally representative samples of stores. These data also furnish important controls including price, availability, and customer characteristics.

We use this approach to estimate the causal effect of American military casualties in the Iraq War on Americans' national identification, the strength of national identity. National identity, like other dimensions of social identity, is a category with which people affiliate ("I am American.") that contains prescriptive behavioral norms that, when followed, foster self-esteem (Tajfel and Turner, 1979; Hogg, 2006; Akerlof and Kranton, 2000). For example, Americans with stronger national identification prefer American brands (Shimp and Sharma, 1987).

War casualties represent external threats to the nation. A core element of current nationalist political strategy is to emphasize and exaggerate external threats to the nation like war, trade competition, and immigration (Mutz, 2018). In social identity theory, external threats to the group threaten the self-esteem derived from group membership (Sedikides, 1993; Sedikides and Strube, 1997). The typical psychological response to external threats is to strengthen group identification (Branscombe et al., 1999; Davies et al., 2008), making emphasis of external threat a convenient nationalist political strategy. Relevant for casualties, images of death in particular generate this response (Greenberg et al., 1994).

We estimate the causal effect of local Iraq War casualties on change in local market share of American-sounding brands during 2003-2006. For a given supermarket, a local Iraq War casualty is an American soldier who died in Iraq, and whose hometown is in the same US county as the store. Americans perceive strong connections to local war casualties (Gartner, 2009; Kriner and Shen, 2010). Local casualties are particularly suited to isolating the effects of external threats on national identification. Conditional on local military enlistment, local exposure to casualties is quasi-random.

We use weekly supermarket scanner data that covers sales of over 8,000 brands for a representative sample of more than 1,100 US stores during 2001-2006. Brand nationality is a cue that operates outside of consumers' conscious awareness in a manner analogous to stereotypes (Martin et al., 2011). Our perception-based measure of brands' perceived American nationality approximates consumers' reliance on brand names to infer products' country of origin (Samiee et al., 2005) and also captures consumers' existing brands associations that shape priors about brands' country of origin.

Our research design isolates casualties' effects on change in the sales of American-sounding brands. For each week during 2003-2006, we estimate the change in market share of American-sounding brands in that store-week relative to the same store-week in 2001. Change since 2001 provides a pre-war baseline against which to measure change.[1] This design holds constant relatively time-invariant correlates of consumption, including customer demographics, the ex ante demand for American-sounding brands, brand availability, and product characteristics such as quality that often underlie consumers' evaluations of domestic and foreign brands (Gürhan-Canli and Maheswaran, 2000). The design also holds constant ex ante variation in consumers' preference for American-sounding brands. To the extent that Americans with stronger national identification already purchased American-sounding brands, local casualties would not change these brands' market share. Our identifying assumption is that local casualties did not otherwise systematically affect the market share of American-sounding brands. Local casualties did not influence product supply nor did they consistently prompt calls to change consumption behavior. We control for weekly changes in brand price and selection, weekly local economic conditions, county population, and ZIP code-level military

[1] Any 2001 events that strengthened national identification - such as 9/11 and Afghanistan war casualties - bias against our expected finding.

enlistment.

Our baseline results show that, on average, the market share of American-sounding brands increased in store-weeks exposed to local causalities. This results holds for each year during 2003-2006. Among the most American-sounding brands, the casualty-induced rise in market share is roughly half the effect of a two-standard-deviation price decrease. Our baseline finding is robust to multiple plausible measures of local casualty exposure and the omission of high-casualty outlier counties. We include additional controls that could mediate casualty-induced change in national identification and/or brand choice, including local partisanship and weekly national casualty counts, which account for national-level factors like events in Iraq, national media coverage, and elite rhetoric. Additionally, we show that local casualties' effects are not fleeting, but endure over our four-year sample.[2] We find that the rise in market share from a one standard deviation increase in cumulative casualties is roughly one third the effect of a two-standard-deviation drop in price.

We further unpack casualties' effects on national identification. We assess whether strengthened national identification requires opposition to a distinct out-group. Our brand sample includes products linked to "Coalition of the Willing" countries, US allies whose troops fought along side Americans in Iraq, and vocal opponents like France and Germany. Neither set of countries were direct external threats but varied in their support of US interests. We find that market shares for brands associated with both sets of countries declined by similar magnitudes throughout the sample period. The finding is consistent with strengthened national identification without focused animosity towards an out-group. Further, it shows that Americans do not construe allied countries to be part of a larger in-group.

We also evaluate demographic heterogeneity in response to local casualties. Our research design holds constant ex ante heterogeneity but customer demographics can mediate the relative salience of national identity and the stickiness of brand preferences. Stores with higher proportions of US citizen customers and relatively low customer educational attain-

[2]Weekly scanner data cannot generate accurate inferences about the persistence of shocks into future weeks. Lasting effects of the casualties cannot be parsed from unobserved shocks in subsequent weeks.

ment see, on average, larger increases in American-sounding brands' market share. Market shares exhibit smaller gains post casualty in stores with greater customer racial and ethnic diversity. Finally, we find suggestive evidence that Iraq War-related elite cues alone do not drive strengthened identification but may magnify local casualties' effects.

Our study demonstrates that external threat strengthens national identification enough to systematically influence choice behavior in real time and into the future. This approach is a powerful complement to experiments, which can more readily analyze dimensions like emotion and information processing, and national identities' substantive content (Huddy and Khatib, 2007; Abdelal et al., 2009; Barabas and Jerit, 2010). The approach offers a compelling alternative to convenience samples (Hafner-Burton et al., 2017).

Observed consumption is a particularly stringent test of change in identification; we show that external threats prompt Americans to spend their own money to reinforce their national identity in everyday behavior. Whereas standard behavioral outcomes like voter turnout analyze selection into a behavior (Kriner and Shen, 2010; Koch and Nicholson, 2016), we hold constant selection and evaluate the content of behavior. By analyzing the political drivers of consumption, we also gain broader insights into the political behavior of those who are unable or unwilling to engage in more direct forms of political participation.[3] To the extent that disparities in direct political participation fuel fundamental shifts in political polarization and wealth inequality, rigorous analysis of indirect participation can help identify and explain disparities.

Our findings also unpack the foundations of public responses to war. Whereas most extant research treats casualties as a type of information, we emphasize casualties' consequences for social identity.[4] Current findings are mixed. Casualties may reduce public support if, within a cost-benefit analysis, they indicate a lower probability of winning (Gartner and Segura, 1998; Gartner, 2008); increase support by strengthening resolve (Gelpi et al., 2009); or have

[3]This mechanism is qualitatively different from explicitly political consumption such as consumer boycotts, which are typically organized and targeted towards specific products/producers (Kam and Deichert, 2019).

[4]Althaus and Coe (2011) and Koch and Nicholson (2016) are exceptions.

no consistent effect because the public is largely unaware of casualties, especially early in conflicts (Groeling and Baum, 2008; Berinsky, 2009).

Our findings indicate that Americans are, in fact, aware of war casualties, but national identification does not drive levels of expressed support for war. On average, Americans were aware of and responsive to local casualties throughout the sample period, despite variation in media coverage and elite cues. Whereas existing work emphasizes the absence of information/cues, our findings point to differences in the assessment of information. National identification and war support diverged. Figure 1 demonstrates divided war support nationally in the second half of the sample. War support tended to decline in casualty-affected communities specifically (Althaus et al., 2012; Kriner and Shen, 2012). The divergence indicates that national identification does not translate into greater acceptance of policy/leaders' actions (Federico et al., 2005; Althaus and Coe, 2011). Contrary to experimental findings (Levendusky, 2017), strengthened national identification did not ease partisan polarization.

Our findings also offer broader insights into the contemporary evolution of American nationalism. We return to them in the conclusion.

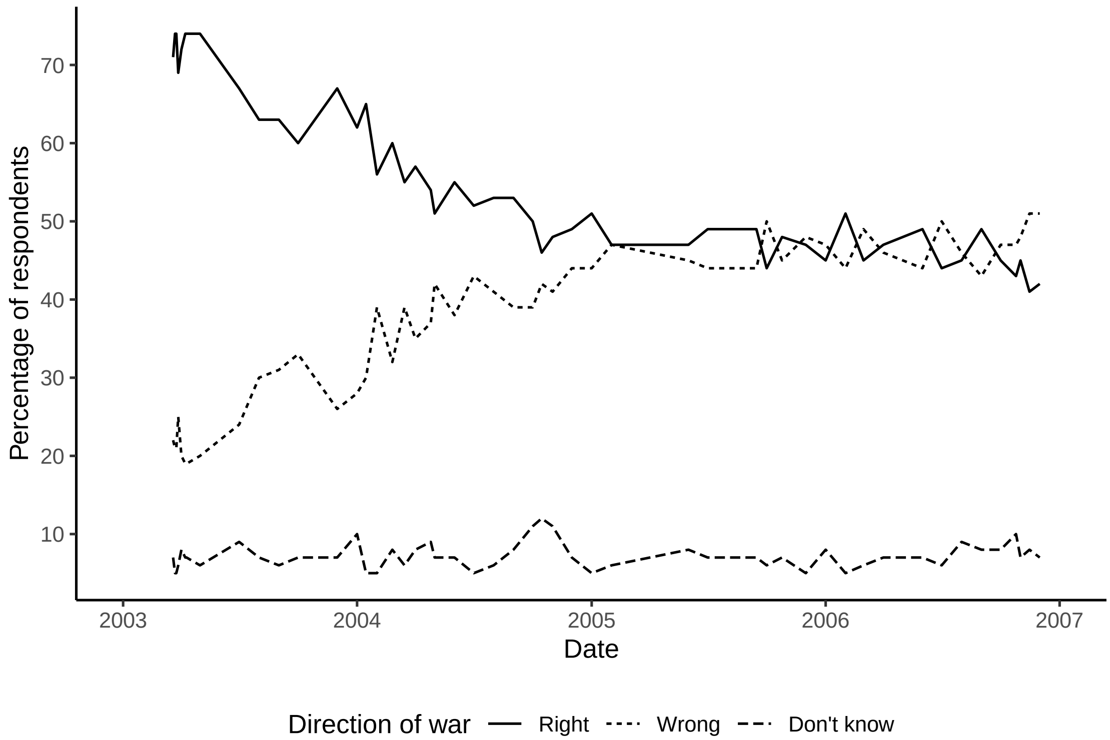
2 Background: Local Exposure to Iraq War Casualties, 2003-2006

The 2003 Iraq War generated 3,240 U.S. military casualties during 2003-2006, about 90 percent of total US war casualties during these years. Figure 2 plots weekly Iraq casualty counts. The weekly average is relative stable. The largest spikes correspond to predictable moments, including the initial invasion and the expanding insurgency in 2004.

From the perspective of a given community, a "local" casualty refers to the death of a deployed soldier originally from that community. Approximately 2,000 unique American cities experienced at least one war casualty during the sample period; 18 cities had ten or more casualties.[5] Appendix Figure A.1 maps cumulative county-level American causalities

[5]See Appendix Tables A.1, A.2, and A.3 for additional details on the distribution of casualties across cities, states, and service units.

Figure 1: American Support for Iraq War Declined, 2003-2006



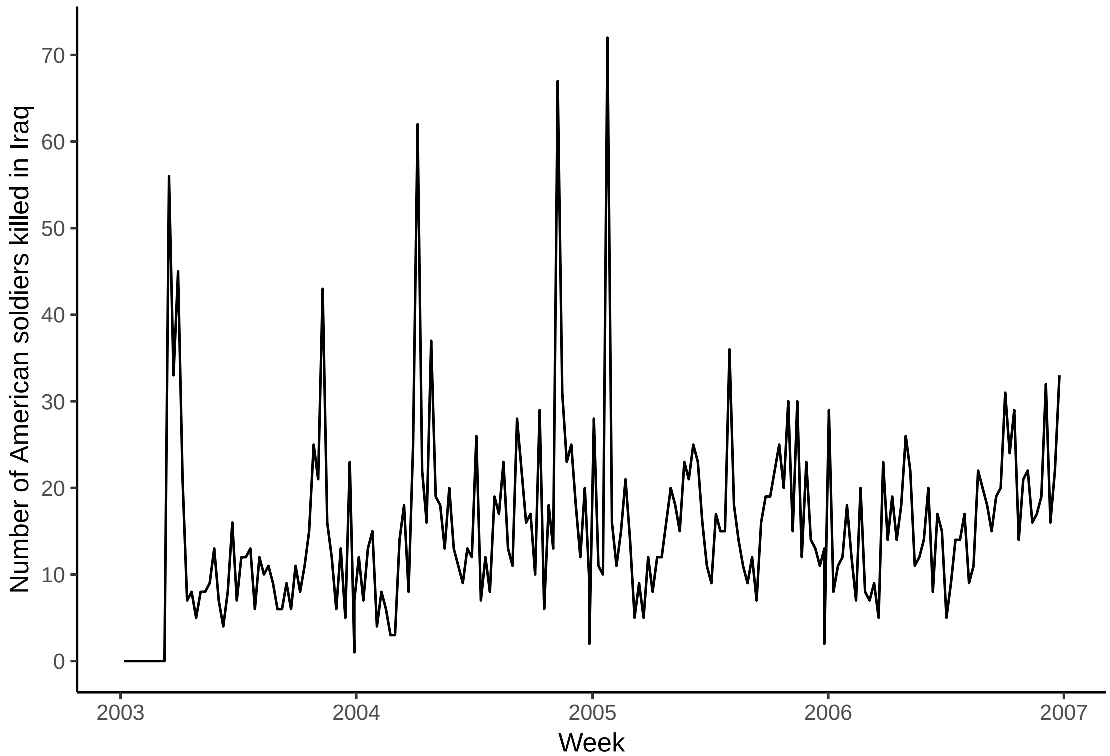
Data Source: Pew Research Center.

in Iraq during 2003-2006. Casualties' geographic distribution is roughly proportional to population density.

Public support for the Iraq War shifted dramatically during this period. Figure 1, which draws on polling data from Pew Research Center, shows that initially more than 70 percent of respondents believed that the war was headed in the right direction. Sharp divisions emerged such that by the 2006 midterm elections, a slight majority of Americans disapproved of the war. This shift in public support is an insightful backdrop for our analysis of national identification. As we analyze shifts in national identification during the same time period, we can draw broad generalizations about the correlation between identification and support for policy.

Our identifying assumption is that conditional on military enlistment, a community's exposure to war casualties is quasi-random. Timing of local casualties is clearly exogenous

Figure 2: Weekly Iraq War Casualties, 2003-2006



Data Source: Associated Press.

to any given American community.[6] Military enlistment, however, is non-random, because during the Iraq War, the US had an all-volunteer military. To the extent that communities with higher enlistment are systematically different from other communities, we can draw sound inferences about only those communities.

An additional assumption is that consumers are aware of local casualties. Proximity increases the likelihood of exposure to information about the casualty.[7] Local casualties produce the largest shifts in war support among non-consumers of news media, suggesting social networks, rather than media, transmit this information (Althaus et al., 2012; Kriner and Shen, 2012). The large fraction of survey respondents who reported knowing a solider who died in Iraq - implausibly high given the actual number of deaths - suggests a propensity to perceive connection to casualties (Gartner, 2009). With the exception of gender, demo-

[6]Soldiers' hometowns are distinct from the location of their service units. See appendix for distribution of casualties by service unit.

[7]We follow existing research in defining local casualties at the county level.

graphic characteristics do not correlate with self-reported connection to casualties (Gartner, 2009). Self-reported connection increased the likelihood that the Iraq War drove choice of political candidate (Gartner, 2009) and lowered presidential approval ratings (Gartner, 2008).

2.1 Local War Casualties as External Threats

We argue that local casualties are external threats that strengthen national identification. Casualties can represent a threat to American values for which soldiers sacrificed their lives (Branscombe et al., 1999). More broadly, the psychological theory of mortality salience holds that death-related thoughts prompt more vigorous defense of worldview to buffer against anxiety (Greenberg et al., 1992, 1997; Arndt et al., 1997). War causalities generate vivid and symbol-laden images of death such flag-draped coffins (Gartner, 2011).

This defense magnifies in-group solidarity. Scholars of mortality salience have shown strengthened national identification to be one of the most consistent responses to death-related cognition (Jonas et al., 2002), including more favorable evaluations of statements that praise the US (Greenberg et al., 1994) and opposition to the desecration of cultural symbols like the American flag (Greenberg et al., 1995). Of direct relevance to this study, mortality salience consistently increases demand for own-country products. Primed to contemplate death, experimental subjects more strongly prefer domestic brands of soda and chocolate (Friese and Hofmann, 2008), and automobiles (Nelson et al., 1997). Patriotic consumers are more likely to switch to domestic brands following death-related media coverage (Liu and Smeesters, 2010). Koch and Nicholson (2016) argue that war casualties increased voter turnout due to mortality salience.

We note other possible meanings of casualties. They can be reminders of personal safety threats from terrorism, the war's stated motive. This is a variant of our proposed mechanism. Another possibility, in line with extant research on casualties and war support, is that casualties provide information about the prospects of winning of the war or the country's performance in waging the war. That information could plausibly influence national identification in connection with strengthened resolve, for example. Though we cannot rule this out completely, our baseline finding shows strengthened identification is consistent throughout the war, whereas we might expect this mechanism to produce variation in response in accordance with war performance.

3 Data and Measurement

Our empirical analysis requires measurement of three central concepts: weekly supermarket purchases, the perceived American origin of supermarket brands, and local exposure to Iraq war casualties.

3.1 Weekly Supermarket Purchases

We measure consumer response to casualties in Iraq using weekly supermarket scanner data supplied by Information Resources Inc. (IRI), a leading source of such data (Bronnenberg et al., 2008). These data cover a representative sample of 1,145 supermarkets across 50 IRI-designated geographic markets.[8] Figure A.2 maps the geographic coverage of our data. The 135 supermarket chains represented in our sample collectively account for around 80 percent of US supermarket sales during the sample period.[9] Over 90 percent of Iraq War casualties during 2003-2006 were from hometowns covered by these data.

We construct our store-level measures of consumer response using weekly unit sales for 8,644 brands across 28 categories of supermarket products.[10] Major supermarket chains stock mature brands and maintain a relatively stable portfolio of brands within each store.

[8]IRI set its market definitions in 1987 to achieve a representative sample of US consumers, making it unlikely that our findings are an artifact of sample selection. See Appendix Table A.5 for a list of IRI geographic market names.

[9]During this period, supermarkets accounted for roughly 70 percent of all U.S. grocery purchases.

[10]The 28 categories are: beer, blades, carbonated beverages, cigarettes, coffee, cold cereal, deodorant, diapers, facial tissue, frozen dinners, frozen pizza, household cleaners, hot dogs, laundry detergent, butter, mayonnaise, milk, mustard/ketchup, paper towel, peanut butter, razors, salty snacks, shampoo, soup, spaghetti sauce, sugar substitutes, and yogurt.

We aggregate data across multiple stock keeping unit (SKU) codes of a single brand-product category, but not across distinct, but related, brands.[11] In addition to unit sales, the IRI data reports price and the number of product variants, which we use as control variables.

3.2 Perceived Brand Nationality

We measure perceived brand nationality using a product's brand name because it is a highly salient, readily available cue for consumers (Usunier and Shaner 2002).[12] For American consumers, brand names based on foreign languages frequently cue associations with a foreign country through distinctive letter combinations and special characters (like umlauts and accent marks) that do not occur in English. By contrast, brands that incorporate geographic locations in the US or American cultural symbols imply American-made products. Survey and experimental evidence indicate that consumers systematically misidentify the national origin of products because they infer nationality from marketing cues, rather than searching for official country-of-origin labels (Samiee et al., 2005; Balabanis and Diamantopoulos, 2011).[13] Consumers draw inferences about nationality based on prior associations between the implied country of origin and the product. A French-sounding brand name, for instance, cues "a rich network of associations related to aesthetic sensitivity, refined taste, and sensory pleasure" (LeClerc et al 1994, 264-268).

To assess the perceived nationality of brands by American grocery shoppers, we administered surveys via Amazon's Mechanical Turk (MTurk). Our survey presented respondents with a product's brand name and category, and asked them to select the most relevant from a list of brand nationalities.[14] Across a range of disciplines, MTurk results are often

[11] For instance, six-packs and two-liter bottles of Coca-Cola are distinct SKUs within the same brand. Coca-Cola and Diet Coke are separate brands.

[12] We performed a trial experiment to test whether additional brand information influenced perceived nationality. For a random sample of brands with US-trademarked logos, we surveyed a randomly selected group on the nationality of brands based on the brand name, product category, and logo. A control group scored the same brands based solely on brand name and product category. Responses were not statistically distinguishable between the two groups.

[13] Products labeled "made in the USA" must meet legal requirements set by the US Federal Trade Commission.

[14] We conducted our survey in 2011. While it is possible that survey captures changes in marketing

Table 1: Brand Examples Across $\mathbf{AmericanScore_i}$ Values

$AmericanScore_i$ Values	Brand Example (Product Category)
7	Sam Adams Boston Lager (beer) Kentucky Gold (ketchup/mustard)
6	Land O' Lakes (margarine/butter) Phillies (hot dogs)
5	Olde Cape Cod (spaghetti sauce) Swanson American Recipes (frozen dinners)
4	New England (ketchup/mustard) Dad's Root Beer (carb. beverages)
3	Maple Leaf (hot dogs) Van De Kamps (frozen dinners)
2	Life in Provence Aioli (mayonnaise) Dietz & Watson (ketchup/mustard)
1	Royal Scot (margarine/butter) World Trend (toothbrushes)
0	König Ludwig Weiss (beer) Cucina Antica (spaghetti sauce)

$AmericanScore_i$ = Number of survey respondents that deem brand b to be American

more reliable results than convenience- and lab-based samples (Casler et al., 2013; Berinsky et al., 2012; Huff and Tingley, 2015; Erlewine and Kotek, 2016). While MTurk samples may not mirror American demographics (Levay et al., 2016), the gap is likely uncorrelated with product nationality perceptions.

Our primary measure of perceived product nationality, $AmericanScore_i$, takes values 0-7, corresponding to the number of respondents who deemed brand b to be American. Table 1 provides examples of brands at each value of $AmericanScore_i$. Brands with $AmericanScore_i =$ 7 exhibit strong nationality cues, including geographical references and references to historical American figures. Low-scoring brands have distinctive foreign elements, including non-English words and foreign geographic references.[15]

strategies after the sample period, core branding features of mature brands are highly stable.

[15] Appendix Table A.4 illustrates the distribution of American-sounding brands across the 28 product categories in our sample.

3.3 Exposure to U.S. War Casualties

We identify weekly local American casualties in Iraq using data from official US Department of Defense press releases compiled by the Associated Press. For each casualty, we have name, hometown, rank, unit, and date and cause of death.[16] We match each casualty to the county in which their hometown is located. We measure casualties at the county level because it is the most conservative measure of exposure we can accurately construct and is consistent with existing studies of local casualty responses.

We use these data to create an indicator variable equal to one if a casualty occurred in the same county as store j in week t, and zero otherwise. 92 percent of store-weeks exposed to casualties experienced a single death. [17] We verify our results are robust to a continuous measures of local weekly casualty exposure.

3.4 Additional Control Variables

Our identifying assumption of quasi-random casualty exposure is conditional on military enlistment. We control for the sum of enlistment over the previous five years for the ZIP code in which store j is located. These data are based on enlistees' home addresses and cover all military branches.[18] Controlling for enlistment also accounts for unobserved local characteristics correlated with both military enlistment and the propensity to react to local casualties by changing consumption.

Changes in local economic conditions are highly unlikely to correlate with casualty exposure and our sample precedes the economic decline that culminated in the Great Recession. We nonetheless control for average home prices in store j's ZIP code and week t. Fluctua-

[16] We verify that hometown is distinct from the location of the casualty's service unit. For example, among the 294 casualties of soldiers based at Fort Hood, Texas, there are 259 hometowns across 51 US states and territories.

[17] An additional 7 percent of store-weeks in which a casualty occurred represented two casualties. The maximum number of casualties in a given store-week during this time period was four casualties, which represented less than 0.5 percent of total store-weeks with the occurrence of a casualty.

[18] Data are from Kriner and Shen (2010) and originally obtained through a Freedom of Information Act request to the U.S. Department of Defense.

tions in local economic conditions may influence the purchase of American-sounding brands through unobserved mechanisms such as changes in the propensity to learn about casualties or perceive external threats.[19]

Partisanship may systematically influence the propensity to respond to external threats by changing consumption. Conservatives tend to have stickier brand preferences because they value consistency and stability more so than liberals (Khan et al., 2013; Jost et al., 2017). We control for this using county vote share for George W. Bush in 2000. Data are from Dave Leip's Electoral Atlas.

Finally, county population size may influence the propensity to know about and respond to local casualties. We also control for county population in 2000. Data are from the 2000 U.S. decennial census.

4 Empirical Analysis

For each product-category store-week in our sample, we model the change in market share growth between 2001 (prior to the beginning of the Iraq war) and years 2003-2006 at each level of $AmericanScore_i$. Our outcome of interest is indexed by:

i: 8 $AmericanScore_i$ levels (0-7),

j: 1,154 supermarkets,

k: 28 product categories, and

t: 52 weeks.

A brand's weekly store market share is the number of brand product units sold as a percentage of all units in the product category sold in that store-week. For example, if brand b in product category k (e.g., yogurt) had a 50% market share in a given store j for week t, the brand accounted for half of all units of yogurt sold in that store in that week.

[19] We verify that American-sounding brands are not systematically less expensive such that tighter budget constraints prompt a switch to lower cost brands.

Measuring market share, as opposed to the total number of units sold, allows us to scale that store's sales of a brand relative to overall demand for that product category in that store-week. Changes in market share also capture shifts in demand for brands distinct from changes in demand for a particular product category. For each product-category store-week, we calculate the average change in market share across brands at each of the eight levels of $AmericanScore_i$. This aggregation reflects our interest in change across $AmericanScore_i$ levels rather than individual brands and reduces the sample to a computationally feasible size. As compared to sampling a subset of stores, this approach minimizes computational burden, maintains generalizability, and utilizes variation in casualties across all stores.

Within each year during 2003-2006, for every $AmericanScore_i$ level-product category-store-week in our sample, we calculate the change in market share between week t in that year and same week in 2001($ShareYear\text{-}2001_{ijkt}$). Measuring change in demand within each store allows us to hold constant all time-invariant baseline characteristics of the store's customer base that influence sales, including ex ante customer preferences. It also accounts for seasonal fluctuations such those due to patriotic holidays like July 4th. We choose 2001, the first year for which scanner data are available, as a baseline because it precedes almost all war casualties.[20] For each store, we retain only brands that were sold in all weeks of the given year and 2001, so our results are not biased by attrition and entry.

We estimate a difference-in-differences ordinary least squares model of weekly changes in market-share growth ($\Delta ShareYear\text{-}2001_{ijkt}$):

$$\begin{aligned}\Delta ShareYear\text{-}2001_{ijkt} = {} & \beta_1 LocalCasualty_{jt} + \beta_2 AmericanScore_i + \beta_3 LocalCasualty_{jt} * AmericanScore_i + \\ & \beta_4 Enlistment_j + \beta_5 Enlistment_j * AmericanScore_i + \beta_6 HomePrice_{jt} + \\ & \beta_7 HomePrice_{jt} * AmericanScore_i + \beta_8 Population_{j2000} + \\ & \beta_9 \Delta PriceYear\text{-}2001_{ijkt} + \beta_{10} \Delta VariantsYear\text{-}2001_{ijkt} + \epsilon_{ijkt}\end{aligned}$$

[20] If 9/11 and/or the eleven US war casualties in 2001 increased sales of American-sounding brands, this would bias against our expected finding for subsequent years.

where

$$
\begin{array}{lcl}
Year & \in & [2003, 2004, 2005, 2006], \\
\Delta \mathit{ShareYear\text{-}2001}_{ijkt} & = & \text{average difference in market share from 2001 to } \mathit{Year} \text{ for} \\
 & & \mathit{AmericanScore}_i \text{ in store } j \text{ belonging to product} \\
 & & \text{category } k \text{ at week } t, \\
\mathit{LocalCasualty}_{jt} & = & \text{indicator for US war casualty in Iraq from same county} \\
 & & \text{as store } j \text{ in week } t, \\
\mathit{AmericanScore}_i & = & \text{scale from 0-7 indicating level of perceived American origin,} \\
\mathit{Enlistment}_j & = & \text{total military enlistment in same ZIP code as store } j \\
 & & \text{in last five years,} \\
\mathit{HomePrice}_{jt} & = & \text{average home price in same ZIP code as store } j \text{ and week } t, \\
\mathit{Population}_{j2000} & = & \text{population in same county as store } j \text{ in year 2000,}
\end{array}
$$

$$
\begin{array}{lcl}
\Delta \mathit{PriceYear\text{-}2001}_{ijkt} & = & \text{average difference in price from 2001 to } \mathit{Year} \text{ for} \\
 & & \mathit{Americanscore}_i \text{ in store } j \text{ belonging to product} \\
 & & \text{category } k \text{ in week } t, \\
\Delta \mathit{VariantsYear\text{-}2001}_{ijkt} & = & \text{average difference in number of variants from 2001 to } \mathit{Year} \\
 & & \text{for } \mathit{AmericanScore}_i \text{ in store } j \text{ belonging to product} \\
 & & \text{category } k \text{ in week } t, \text{ and} \\
\epsilon_{ijkt} & = & \text{normally distributed random error term.}
\end{array}
$$

The coefficient of interest is β_3, on the interaction between local casualty exposure and $\mathit{AmericanScore}_i$. We estimate separate models for each year in our data (2003-2006). As is standard in empirical marketing analyses, we control for two time-varying brand-store characteristics that influence fluctuations in market share (Ataman et al., 2010). $\Delta \mathit{PriceYear\text{-}2001}_{ijkt}$ controls for exogenous price changes and the effect of promotional, limited-time discounts.[21] Non-price responses, such as advertising, are less likely because they require longer lead times to implement. Price promotions are retailers' fastest response to negative demand shocks.[22] Retailers' contracts with manufacturers forbid changes to products' shelf space allocation and location, so systematic, retailer-driven change in product supply or location is unlikely.[23] We also control for weekly changes in the number of varieties of a brand a store stocks in a product category, $\Delta \mathit{VariantsYear\text{-}2001}_{ijkt}$. All else equal, consumers are more likely to purchase a brand if a store stocks more varieties.

4.1 Baseline Results

Table 2 presents our baseline results for 2003, the first year of the Iraq War. Model (1) focuses on local weekly casualties. The coefficient of interest is $\mathit{LocalCasualty}_{jt} * \mathit{AmericanScore}_i$, the interaction between local casualty exposure and American score. This interaction is positive and statistically significant in both models, indicating that, on average, in weeks that stores experience a casualty, the market share of American-sounding brands grows. Model (2) adds controls for the national casualty environment, the total number of American casualties in Iraq in week t. This variable captures time-varying national factors that could influence perception of external threats including events in Iraq, media coverage, and elite rhetoric. Our finding is unchanged.

Figure 3 plots in black the coefficient on $\mathit{LocalCasualty}_{jt} * \mathit{AmericanScore}_i$ for Model (2), estimated annually during 2003-2006, extracted from our yearly models.[24] The figure illustrates that despite large variation in public and elite opinion towards the war during these years, Americans consistently responded to local casualties by switching to American-sounding brands. On average, the rise in the increase in market share following a local casualty is about half the effect of a two-standard-deviation price decrease for brands per-

[21] We verify weekly price changes are uncorrelated with brands' $\mathit{AmericanScore}_i$.

[22] Manufacturers provide retailers with a trade allowance to finance promotions.

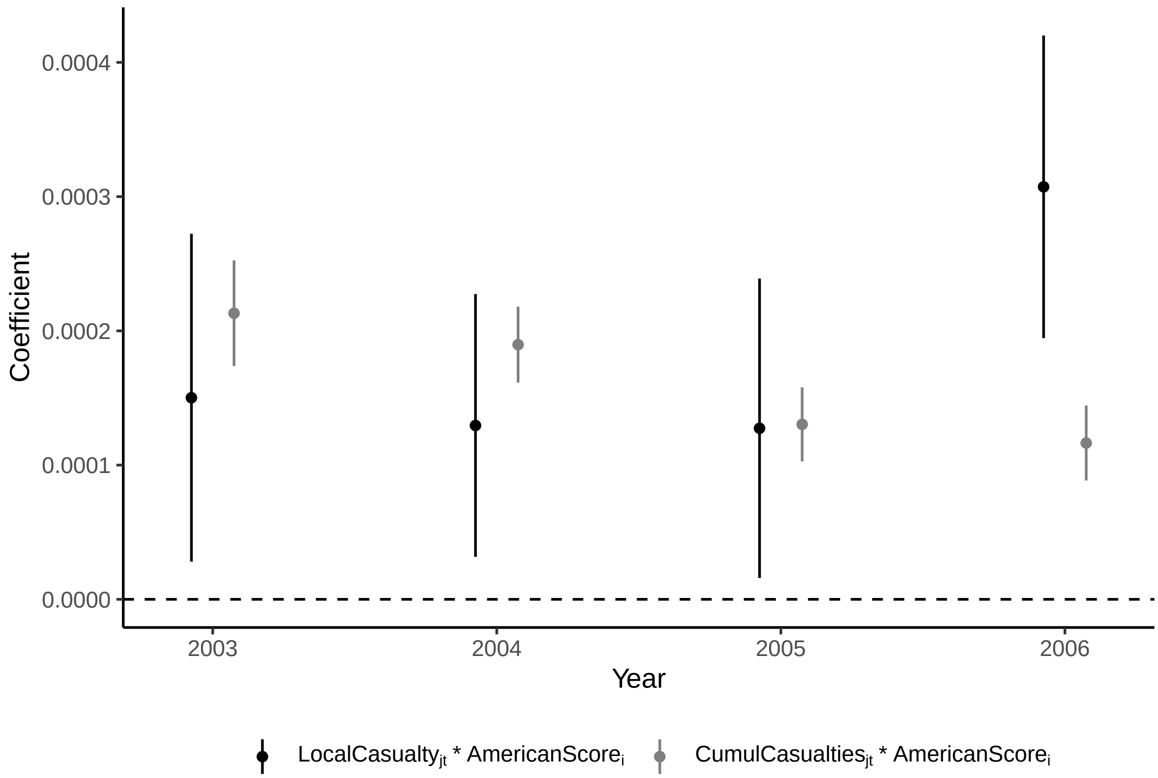
[23] Manufacturers negotiate with retailers for specific shelf locations for their products. Local distributors stock shelves and can monitor compliance. These agreements are negotiated chain-wide and renegotiated at fixed intervals.

[24] See Appendix Table A.6 for underlying model estimates.

Table 2: Weekly Casualties and American Brand Share - 2003

	$\Delta Share2003\text{-}2001_{ijkt}$	
	Local casualties	Local and nat. casualties
	(1)	(2)
$LocalCasualty_{jt} * AmericanScore_i$	0.00016***	0.00014**
	(0.00006)	(0.00006)
$NationalCasualties_t * AmericanScore_i$		0.000001
		(0.000001)
$LocalCasualty_{jt}$	0.00032	0.00025
	(0.00028)	(0.00029)
$NationalCasualties_t$		0.00001
		(0.000004)
$AmericanScore_i$	−0.00012***	−0.00013***
	(0.00002)	(0.00003)
$HomePrice_{jt}$	0.00004***	0.00004***
	(0.000003)	(0.000003)
$HomePrice_{jt} * AmericanScore_i$	−0.00001***	−0.00001***
	(0.000001)	(0.000001)
$Enlistment_j$	−0.00104***	−0.00105***
	(0.00027)	(0.00027)
$Enlistment_j * AmericanScore_i$	−0.00013**	−0.00013**
	(0.00006)	(0.00006)
$Population_{j2000}$	0.0000002***	0.0000002***
	(0.00000)	(0.00000)
$\Delta Price2003\text{-}2001_{ijkt}$	−0.00053***	−0.00053***
	(0.00001)	(0.00001)
$\Delta Variants2003\text{-}2001_{ijkt}$	0.00900***	0.00900***
	(0.00001)	(0.00001)
$Intercept$	−0.00036***	−0.00042***
	(0.00011)	(0.00012)
Observations	6,715,772	6,715,772
R^2	0.10796	0.10797
Note:		*p<0.1; **p<0.05; ***p<0.01

Figure 3: Casualties Increase American Brand Market Share Growth, 2003-2006



Effect of weekly and cumulative casualty exposure on market share of American-sounding brands over time, with 95 percent confidence intervals. Coefficients for weekly casualties drawn from Appendix Table A.6. Coefficients for cumulative casualties drawn from Appendix Table A.7.

ceived as most American ($AmericanScore_i = 7$).

We submit our baseline results to robustness checks.[25] We lag casualties by one week to allow for a delay in consumer response; our results are largely unchanged. Our findings are also unchanged if casualty exposure is based on American casualties in both Iraq and Afghanistan. We find no effect of just Afghanistan war casualties, likely because they accounted for just ten percent of U.S. military causalities during the sample period. Political conservatives tend to have stickier brand preferences (Khan et al., 2013). Our results do not change when we add George W. Bush's 2000 vote share for the county in which store j is located.

[25] All results available upon request.

4.2 Cumulative Casualties

Our baseline results demonstrate that consumers respond to local casualty exposure by purchasing more American-sounding brands at the supermarket. However, they grant little insight into whether this effect is short-lived, or if it accumulates over time as a result of repeated exposure to wartime casualties. We answer this question by leveraging the same empirical strategy, but replacing weekly casualties with the natural log of cumulative local casualties in the same county as store j from the beginning of the war to week t. We also control for the natural log of national cumulative casualties, in line with our baseline specifications.

Table 3 presents our results for cumulative casualties for 2006, the last year in our sample.[26] Model (1) focuses on logged local cumulative casualties, while Model (2) again adds controls for the national cumulative casualty environment. In line with our baseline results, repeated exposure to casualties over time increases the change in market share for more American-sounding brands. A one standard deviation increase in cumulative casualties results in an increased change in market share that is approximately one third of that associated with a two-standard-deviation drop in price for brands perceived as most American ($AmericanScore_i = 7$). The effect remains when we take into account the national cumulative casualty environment, indicating that the cumulative effect of casualty exposure on national identification operates independent of the war's trajectory and total costs, and changes in elite rhetoric and access to information.

Figure 3 displays in gray the coefficient on the interaction between logged local cumulative casualties and $AmericanScore_i$, and its 95 percent confidence interval, for every year from 2003 to 2006, extracted from our yearly models controlling for national cumulative casualties. While the estimated coefficient decreases in size over time, note that this also coincides with a monotonically increasing number of local cumulative casualties over time.

[26] See Appendix Table A.7 for full results for all years 2003-2006.

Table 3: Cumulative Casualties and American Brand Share - 2006

	$\Delta Share2006\text{-}2001_{ijkt}$	
	Local casualties	Local and nat. casualties
	(1)	(2)
$ln(CumulCasualties_{jt}) * AmericanScore_i$	0.00011***	0.00010***
	(0.00001)	(0.00001)
$ln(CumulNatCasualties_{jt}) * AmericanScore_i$		0.00090***
		(0.00015)
$ln(CumulCasualties_{jt})$	−0.00044***	−0.00060***
	(0.00007)	(0.00007)
$ln(CumulNatCasualties_{jt})$		0.00724***
		(0.00066)
$AmericanScore_i$	−0.00099***	−0.00809***
	(0.00003)	(0.00115)
$HomePrice_{jt}$	−0.00002***	−0.00002***
	(0.000003)	(0.000003)
$HomePrice_{jt} * AmericanScore_i$	0.000003***	0.000003***
	(0.000001)	(0.000001)
$Enlistment_j$	−0.00269***	−0.00255***
	(0.00035)	(0.00035)
$Enlistment_j * AmericanScore_i$	0.00040***	0.00042***
	(0.00008)	(0.00008)
$Population_{j2000}$	0.0000002***	0.0000002***
	(0.0000000)	(0.0000000)
$\Delta Price2006\text{-}2001_{ijkt}$	−0.00019***	−0.00019***
	(0.00001)	(0.00001)
$\Delta Variants2006\text{-}2001_{ijkt}$	0.00968***	0.00968***
	(0.00001)	(0.00001)
Intercept	0.00472***	−0.05248***
	(0.00014)	(0.00524)
Observations	5,533,301	5,533,301
R^2	0.14027	0.14046
Note:		*p<0.1; **p<0.05; ***p<0.01

We perform multiple robustness checks of our cumulative casualty result.[27] We repeat the robustness checks from our baseline analysis: lagged casualty exposure, Iraq and Afghanistan casualties, Afghanistan casualties only, and Bush 2000 vote share. We also remove from the sample all counties with cumulative casualties more than two standard deviations above or below the mean. Finally, we weight local cumulative casualties by county population in 2000 because more populous counties are likely to have higher cumulative casualties (see Appendix Table A.8).

5 Unpacking External Threat and National Identification

We have demonstrated that local exposure to wartime casualties strengthened national identification in the context of the Iraq War, and that this effect was potentially cumulative. In this section, we unpack these effects to establish more nuanced mechanisms.

5.1 National Identification Versus Out-Group Animosity

Strengthened national identification is consistent with multiple variants of national identity that posit distinct orientations towards out-groups (Huddy and Khatib, 2007). Parsing these distinctions can provide insight into whether strengthened national identification necessarily is in direct opposition to an out-group.

The Iraq War provides a unique opportunity to consider how military alliances shape perceptions of group boundaries. Wartime alliances are thought to forge special relationships between countries that are the basis for enduring affinities among their citizens (Siverson and Emmons, 1991). The war divided America's western European allies. France and Germany opposed the invasion while the United Kingdom, Italy, and Spain sent troops to Iraq as members of the "Coalition of the Willing".

We estimate change in the market share of brands associated with the two sets of countries. If wartime alliances/division shape affinities, casualties would increase the market

[27] Results available upon request.

share of brands associated with Coalition of the Willing countries, while lowering the market share of French- and German-sounding brands. Comparison of these two groups also allows to differentiate between a proactive shift towards American brands from a shift against French and German brands stoked by public calls for Americans to boycott products from these countries (Davis and Meunier, 2011).

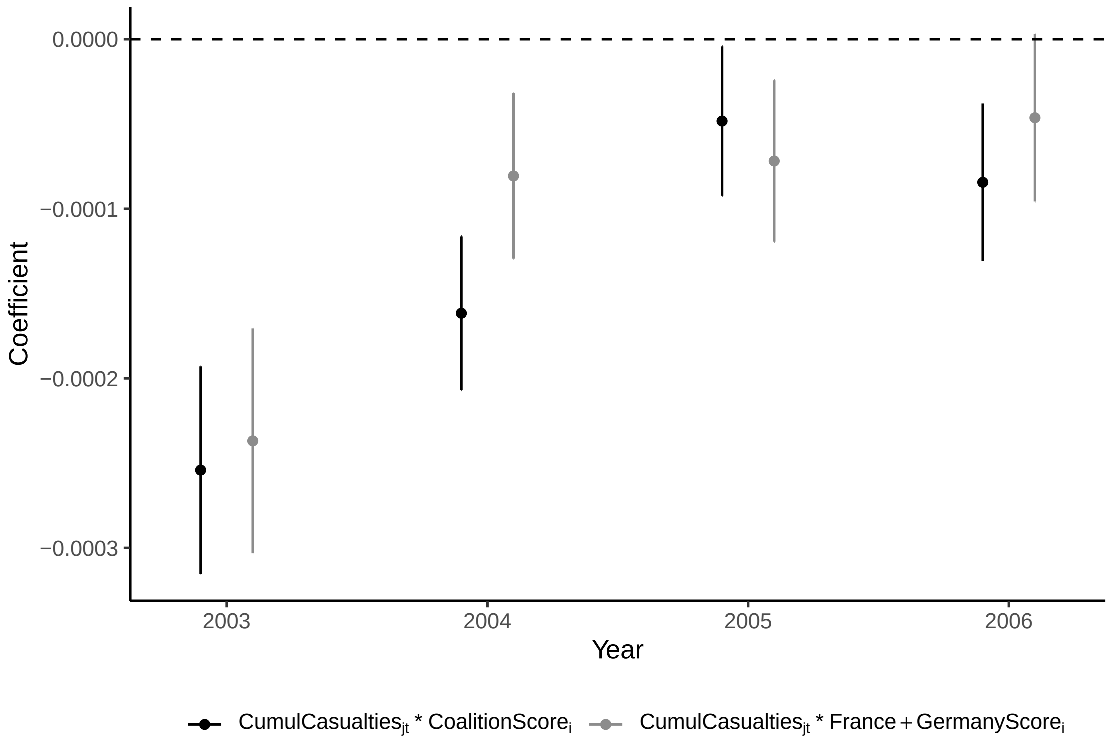
We construct perceived nationality scores from the same original survey data from which we generated $AmericanScore_i$. $CoalitionScore_i$ is equal to how many respondents coded brand b to be from the United Kingdom, Spain, or Italy. $France + GermanyScore_i$ is an analogous measure for France and Germany. Both measures range from 0-7. We estimate two versions of our cumulative causality model, replacing $AmericanScore_i$ with $CoalitionScore_i$ and $France + GermanyScore_i$, respectively. Figure 4 plots the coefficient of the interaction between logged cumulative local casualties and $CoalitionScore_i$ (black) and $France + GermanyScore_i$ (gray) for each sample year.[28] Market shares for both sets of brands declined similarly in response to cumulative casualty exposure, a decline that, for almost all country group-years, is statistically different from zero. For both country groups, the magnitude of market share drop attenuates over time. These findings indicate that military alliances do not engender a sense of shared group identity that manifests in consumers' revealed affinity towards allied countries, and that casualty exposure perhaps only reinforces a sense of American national identity. That we find no distinctive decline in French/German brands suggests that casualties did not sharpen antipathy towards out-groups.

5.2 Demographic Variation

We next consider demographics as a source of heterogeneity in consumer response to external threat. Our difference-in-differences research design holds constant the effects of ex ante heterogeneity in the supply and demand of American-sounding brands. Customer demographics may, however, influence the propensity to respond to local casualties by switching brands.

[28]See Appendix Tables A.9 and A.10 for full results for all years 2003-2006.

Figure 4: Casualties Reduce Market Share of French/German and Coalition of the Willing Brands



Effect of cumulative casualty exposure on market share of "Coalition of the Willing" and French and German brands over time, with 95 percent confidence intervals. Coefficients for "Coalition of the Willing" brands drawn from Appendix Table A.9. Coefficients for French and German brands drawn from Appendix Table A.10.

Existing research demonstrates that demographic factors like race, education, and nativity are meaningful in shaping baseline attachment to an American identity (Schildkraut, 2014). Our empirical setup allows us to investigate if those same demographic factors influence Americans' dynamic responses to external threats. More generally, it allows us to explore how casualties interact with broader demographic cleavages in American politics.

IRI supplies customer demographic data based on the population characteristics for a two-mile radius around each store; these cross-sectional data are derived from the 2000 US Census. We include the percentage of local population that is native-born, along with its square, to test for non-linear effects. Education can shape how individuals respond to external threats to their in-group. We use the percentage of employment that is in blue and white collar jobs to proxy for local level of education.[29] Local armed forced employment accounts for unobserved heterogeneity in local salience of the military. Median household income tests if relatively wealthier areas can more easily respond to war casualties by switching brands. Given that younger individuals may have more malleable psychological predispositions, we include the percentage of the population aged 18-39. The percentage of local population that is Black, Hispanic, and Asian captures possible racial and ethnic variation.

We assess demographic variation in response to cumulative local casualties in 2006, by which point nearly all American casualties in Iraq had occurred. We use a similar modeling strategy to our baseline analyses, but we add a three-way interaction between casualty exposure, perceived brand nationality, and the relevant demographic variables. We run separate models for each demographic variable and verify robustness to excluding stores in counties whose populations are more than two standard deviations away from the mean.

Table 4 presents the results for demographic variables that have statistically significant interactions with cumulative casualties and $AmericanScore_i$ in 2006.[30] In stores with a higher proportion of native-born customers, cumulative local causalities initially magnifies strengthened national identification but magnification is less at higher levels. These results indicate that external threats produce stronger national identification in communities where national identity was ex ante more salient. The findings are consistent with patterns in national identification in response to immigration (Hopkins, 2011; Dancygier, 2010).

Stores with a higher proportion of blue collar worker customers saw relatively higher growth in American-sounding brands' market share in response to local casualties. Market share declined relatively in stores with higher proportions of white collar customers.[31] These findings suggest that increasing educational attainment tends to diminish the in-group-enhancing effect of external threats, like casualties (see Schildkraut, 2014). More educated consumers tend to exhibit a weaker preference for own-country brands (Sharma

[29] As defined by the US Census.
[30] Results for all demographic variables are in Appendix Table A.11.
[31] We also control for household income, which likely correlates with education.

Table 4: Cumulative Casualties and Demographic Variation - 2006

	$\Delta Share2006\text{-}2001_{ijkt}$				
	Native-born	Blue collar	White collar	Black	Asian
	(1)	(2)	(3)	(4)	(5)
$ln(CumulCasualties_{jt}) * AmericanScore_i * NativeBorn_j$	0.00016^{***}				
	(0.00001)				
$ln(CumulCasualties_{jt}) * AmericanScore_i * NativeBorn_j^2$	-0.000001^{***}				
	(0.00000)				
$ln(CumulCasualties_{jt}) * AmericanScore_i * BlueCollar_j$		0.00001^{***}			
		(0.00000)			
$ln(CumulCasualties_{jt}) * AmericanScore_i * WhiteCollar_j$			-0.000008^{***}		
			(0.00000)		
$ln(CumulCasualties_{jt}) * AmericanScore_i * Black_j$				-0.000004^{***}	
				(0.000001)	
$ln(CumulCasualties_{jt}) * AmericanScore_i * Asian_j$					-0.000006^{***}
					(0.000002)
Controls	Yes	Yes	Yes	Yes	Yes
Observations	5,030,086	5,030,086	5,030,086	5,030,086	5,030,086
R^2	0.143	0.143	0.143	0.143	0.143

Note: $^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$

et al., 1995).[32]

Our results with respect to racial and ethnic composition are mixed. The interactions with local share of the population that is black and Asian, respectively, are both negative; the interaction with percentage of local population that is Hispanic is not statistically significant. This broadly suggests that the national identity-enhancing effect of casualty exposure is attenuated (though not eliminated entirely) in places where ethnic and racial minorities are relatively more represented compared to whites.[33] These findings are in line with more general research on minority status and American identity attachment. Individuals who belong to racial and ethnic minority groups are less likely than whites to self-identify as American most of the time, especially people who perceive that their group has historically suffered discrimination (Schildkraut, 2011). Because people belonging to minority groups are less likely to be strongly attached to an American national identity, exposure to casualties does not push areas in which minorities are more represented to switch to American-sounding brands at the same level as areas in which whites are relatively more represented.

5.3 Casualties and Elite Cues

Experimental research emphasizes that elite cues are a key driver of public assessment of war (see, e.g. Berinsky, 2009). We analyze the mediating role of elite cues in local casualty response using data on weekly campaign advertising on the Iraq War during the 2006 midterm elections. We explain how we perform this analysis in detail in Appendix A.1 and display the main associated results in Appendix Table A.12. Overall, the results suggest that greater advertising on Iraq augments the national identification effect of local casualty exposure. This effect appears to be strongest and most consistent in Safe Democratic areas, but also reaches statistical significance at $p < .1$ in Tossup areas. Meanwhile, increased advertising on

[32] Education may also influence reliance on brand cues in product assessments (Evanschitzky and Wunderlich, 2006).

[33] One might be concerned that the race of the casualty is consequential. In previous research on the political impacts of casualties, respondent race does not correlate with reaction to causalities, nor does the race of the casualty produce different effects on the probability supporting war (Gartner and Segura, 2000). Appendix Figure A.3 shows that throughout the Iraq War, casualties are overwhelmingly white soldiers.

Iraq in Safe Republican areas has no detectable interactive effect with cumulative casualty exposure, although as we discuss in Appendix A.1, our coverage of Republican-leaning areas is relatively sparse. When we separately investigate Democratic and Republican advertising, we find evidence that indicates Democratic advertising on the Iraq War largely drives the interactive relationship between casualty exposure and advertising in Safe Democratic and Tossup areas [34] These results suggest that elite cues play a meaningful role in shaping the salience of local casualty exposure.

6 Conclusion

We estimate the causal effect a prominent type of external threat, wartime casualties, on the weekly market share of American-sounding supermarket brands, a real-world behavioral measure of national identification. We show that the market share of these brands increases in stores following the Iraq War death of a solider originally from the same county. Cumulative casualties increase the market share of these brands, indicating that threats are not experienced in isolation but accumulate over time. The shift is a proactive one into American-sounding brands rather than incidental to a shift away from brands associated with other countries. Cross-sectionally, the increase is concentrated in communities with moderate levels of diversity as measured by percent of the area's population that is native-born and that belongs to racial and ethnic minorities, and in communities with lower average educational attainment. We find some evidence that elite priming about the Iraq War magnifies casualties' effect on the market share of American-sounding brands.

Our findings offer reasons for cautious, modest optimism about our current period of nationalist politics. While we show that local casualties consistently strengthened national identification, strengthened identification did not directly translate into support for the Iraq War or incumbent politicians who championed the war (Kriner and Shen, 2010). War is arguably the policy domain most likely to generate unquestioned support of leaders. This

[34] See Appendix Table A.13.

disconnect points to a constructive form of patriotism that combines love of country with constructive criticisms of the country's actions, rather than blind patriotism (Huddy and Khatib, 2007).

We note several possible extensions of this work. Our focus is external threat deriving from war casualties, which represent threats to American values and/or primes thoughts of death. Status threat is a distinct type of threat shown to help explain American vote choices in 2016 (Mutz, 2018). Future research can evaluate the behavioral effects of politically-induced status threat such as through the purchase of luxury goods to bolster self-esteem (Friedman and Sutton, 2013). The emotional underpinnings of threat responses—shame, guilt, rage—likely produce distinctive responses and need further investigation. Cognitive processes are similarly important. For example, Coleman et al. (2019) find that identity threats shapes consumption via effects on memory. Personality traits may also mediate these processes (Feldman and Stenner, 1997).

Another fruitful area of research is explaining the shift in American national identification from the constructive form suggested by strong opposition to the Iraq War towards the blind ethnonationalism form associated with the 2016 US presidential election. While some evidence links Trump 2016 vote share with casualties (Kriner and Shen 2017), our findings suggest that events in the intervening decade produced the shift. Our findings on cumulative casualties point to one possibility: further accumulation of threat after 2006, such as economic distress, which eventually shifted the predominate type of national identification. In this scenario, it may be possible to identify the tipping point after which ethnonationalism took hold. Alternately, elites may capitalized on strengthened identification and engineered the shift towards ethnonationalism. Our campaign advertising findings indicate that elites, though not nationalist political parties per se, can strengthen identification but not create it out of whole cloth. One observable implication might be that during the intervening decade, elites targeted areas with heightened national identification in distinctive ways to shift the content of identification.

More broadly, our research helps specify the psychological foundations of democratic accountability. In establishing that threat responses drive real-time mass behavior, we establish a tighter link between individual responses to social threats and political behavior like protest and voting. Future research can further specify the dimensions of this relationship. For example, shifts in national identification can shape behavior through effects on political trust (Hetherington and Rudolph, 2008). We highlight strengthened national identification as a mechanism that links awareness of international politics to evaluation of leaders. The disconnect between national identification and support for the Iraq War raises important questions about when and how psychological processes translate into expressed attitudes. Comparative research can help map the distinctive sociopolitical foundations of nationalist politics. Key mechanisms are likely different in countries where national identity is more tightly bundled with other social identities like race and religion. Countries whose citizens have weaker baseline attachments to national identity may be less likely to have nationalist political movements, but they may also have fewer dimensions of shared social identity that can anchor democratic accountability.

Finally, our research design can be adapted to other social identities central to political Science, like race and gender, that also feature prominently in product branding. These applications would complement existing political behavior research by testing external validity with the same rigor that surveys establish internal validity.

References

Abdelal, R., Herrera, Y. M., Johnston, A. I., and McDermott, R. (2009). *Measuring Identity: A Guide for Social Scientists.* Cambridge University Press.

Akerlof, G. A. and Kranton, R. E. (2000). Economics and Identity. *Quarterly Journal of Economics*, 115(3):715–753.

Althaus, S. L., Bramlett, B. H., and Gimpel, J. G. (2012). When War Hits Home: The Geography of Military Losses and Support for War in Time and Space. *Journal of Conflict Resolution*, 56(3):382–412.

Althaus, S. L. and Coe, K. (2011). Social Identity Processes and the Dynamics of Public Support for War. *Public Opinion Quarterly*, 75(1):65–88.

Arndt, J., Greenberg, J., Solomon, S., Pyszczynski, T., and Simon, L. (1997). Suppression, Accessibility of Death-Related Thoughts, and Cultural Worldview Defense: Exploring the Psychodynamics of Terror Management, journal = Journal of Personality and Social Psychology, volume = 73, number = 1, pages = 5–18.

Ataman, M. B., Van Heerde, H. J., and Mela, C. F. (2010). The Long-Term Effect of Marketing Strategy on Brand Sales. *Journal of Marketing Research*, 47(5):866–882.

Balabanis, G. and Diamantopoulos, A. (2011). Gains and Losses from the Misperception of Brand Origin: The Role of Brand Strength and Country-of-Origin Image. *Journal of International Marketing*, 19(2):95–116.

Barabas, J. and Jerit, J. (2010). Are Survey Experiments Externally Valid? *American Political Science Review*, 104(2):226–242.

Baum, M. A. (2002). The Constituent Foundations of the Rally-Round-the-Flag Phenomenon. *International Studies Quarterly*, 46(2):263–298.

Berinsky, A. J. (2009). *In Time of War: Understanding American Public Opinion from Wonrld War II to Iraq.* University of Chicago Press.

Berinsky, A. J., Huber, G. A., and Lenz, G. S. (2012). Evaluating Online Labor Markets for Experimental Research: Amazon.com's Mechanical Turk. *Political Analysis*, 20(3):351–368.

Branscombe, N. R., Ellemers, N., Spears, R., and Doosje, B. (1999). The Context and Content of Social Identity Threat. In *Social Identity: Context, Commitment, Content*, pages 35–58. Blackwell Publishers Inc., Oxford, England.

Bronnenberg, B. J., Kruger, M. W., and Mela, C. F. (2008). Database Paper: The IRI Marketing Data Set. *Marketing Science*, 27(4):745–748.

Casler, K., Bickel, L., and Hackett, E. (2013). Separate But Equal? A Comparison of Participants and Data Gathered via Amazon's MTurk, Social Media, and Face-to-Face Behavioral Testing. *Computers in Human Behavior*, 29(6):2156–2160.

Colantone, I. and Stanig, P. (2018). The Trade Origins of Economic Nationalism: Import Competition and Voting Behavior in Western Europe. *American Journal of Political Science*, 62(4):936–953.

Coleman, N. V., Williams, P., and Morales, A. C. (2019). Identity Threats, Compensatory Consumption, and Working Memory Capacity: How Feeling Threatened Leads to Heightened Evaluations of Identity-Relevant Products. *Journal of Consumer Research*, 46(1):99–118.

Dancygier, R. (2010). *Immigration and Conflict in Europe.* Cambridge University Press, Cambridge.

Davies, P. G., Steele, C. M., and Markus, H. R. (2008). A nation challenged: The impact of foreign threat on America's tolerance for diversity. *Journal of Personality and Social Psychology*, 95(2):308–318.

Davis, C. L. and Meunier, S. (2011). Business as Usual? Economic Responses to Political Tensions. *American Journal of Political Science*, 55(3):628–646.

Erlewine, M. Y. and Kotek, H. (2016). A Streamlined Approach to Online Linguistic Surveys. *Natural Language and Linguistic Theory*, 34(2):481–495.

Escalas, J. E. and Bettman, J. R. (2005). Self-Construal, Reference Groups, and Brand Meaning. *Journal of Consumer Research*, 32(3):378–389.

Evanschitzky, H. and Wunderlich, M. (2006). An Examination of Moderator Effects in the Four-Stage Loyalty Model. *Journal of Service Research*, 8(4):330–345.

Federico, C. M., Golec, A., and Dial, J. L. (2005). The Relationship Between the Need for Closure and Support for Military Action Against Iraq: Moderating Effects of National Attachment. *Personality and Social Psychology Bulletin*, 31(5):621–632.

Feldman, S. and Stenner, K. (1997). Perceived Threat and Authoritarianism. *Political Psychology*, 18(4):741–770.

Friedman, R. S. and Sutton, B. (2013). Selling the War? System-Justifying Effects of Commercial Advertising on Civilian Casualty Tolerance. *Political Psychology*, 34(3):351–367.

Friese, M. and Hofmann, W. (2008). What would you have as a last supper? Thoughts about death influence evaluation and consumption of food products. *Journal of Experimental Social Psychology*, 44(5):1388–1394.

Gartner, S. S. (2008). The Multiple Effects of Casualties on Public Support for War: An Experimental Approach. *American Political Science Review*, 102(1):95–106.

Gartner, S. S. (2009). Evaluating Claims of Social Connection to International Conflict Casualties. *International Interactions*, 35(3):352–364.

Gartner, S. S. (2011). On Behalf of a Grateful Nation: Conventionalized Images of Loss and Individual Opinion Change in War. *International Studies Quarterly*, 55(2):545–561.

Gartner, S. S. and Segura, G. M. (1998). War, Casualties, and Public Opinion. *Journal of Conflict Resolution*, 42(3):278–300.

Gartner, S. S. and Segura, G. M. (2000). Race, Casualties, and Opinion in the Vietnam War. *Journal of Politics*, 62(1):115–146.

Gelpi, C., Feaver, P. D., and Reifler, J. (2009). *Paying the Human Costs of War: American Public Opinion and Casualties in Military Conflicts*. Princeton University Press.

Gerber, A. S., Gimpel, J. G., Green, D. P., and Shaw, D. R. (2011). How Large and Long-lasting Are the Persuasive Effects of Televised Campaign Ads? Results from a Randomized Field Experiment. *American Political Science Review*, 105(1):135–150.

Greenberg, J., Porteus, J., Simon, L., Pyszczynski, T., and Solomon, S. (1995). Evidence of a Terror Management Function of Cultural Icons: The Effects of Mortality Salience on the Inappropriate use of Cherished Cultural Symbols. *Personality and Social Psychology Bulletin*, 21(11):1221–1228.

Greenberg, J., Pyszczynski, T., Solomon, S., Simon, L., and Breus, M. (1994). Role of Consciousness and Accessibility of Death-Related Thoughts in Mortality Salience Effects. *Journal of Personality and Social Psychology*, 67(4):627–637.

Greenberg, J., Simon, L., Pyszczynski, T., Solomon, S., and Chatel, D. (1992). Terror management and tolerance: Does mortality salience always intensify negative reactions to others who threaten one's worldview? *Journal of Personality and Social Psychology*, 63(2):212–220.

Greenberg, J., Solomon, S., and Pyszczynski, T. (1997). Terror Management Theory of Self-Esteem and Cultural Worldviews: Empirical Assessments and Conceptual Refinements. In Zanna, M. P., editor, *Advances in Experimental Social Psychology*, volume 29, pages 61–139. Academic Press.

Groeling, T. and Baum, M. A. (2008). Crossing the Water's Edge: Elite Rhetoric, Media Coverage, and the Rally-Round-the-Flag Phenomenon. *The Journal of Politics*, 70(4):1065–1085.

Gürhan-Canli, Z. and Maheswaran, D. (2000). Cultural Variations in Country of Origin Effects. *Journal of Marketing Research*, 37(3):309–317.

Hafner-Burton, E. M., Haggard, S., Lake, D. A., and Victor, D. G. (2017). The Behavioral Revolution and International Relations. *International Organization*, 71(S1):S1–S31.

Hetherington, M. J. and Rudolph, T. J. (2008). Priming, Performance, and the Dynamics of Political Trust. *The Journal of Politics*, 70(2):498–512.

Hogg, M. A. (2006). Social Identity Theory. In Burke, P. J., editor, *Contemporary Social Psychological Theories*, pages 111–136. Stanford University Press, Stanford, CA.

Hopkins, D. J. (2011). National Debates, Local Responses: The Origins of Local Concern about Immigration in Britain and the United States. *British Journal of Political Science*, 41:499–524.

Huddy, L. and Khatib, N. (2007). American Patriotism, National Identity, and Political Involvement. *American Journal of Political Science*, 51(1):63–77.

Huff, C. and Tingley, D. (2015). "Who Are These People? Evaluating the Demographic Characteristics and Political Preferences of MTurk Survey Respondents. *Research and Politics*, 2(3):1–12.

Inglehart, R. F. and Norris, P. (2016). Trump, Brexit, and the Rise of Populism: Economic Have-Nots and Cultural Backlash. Working Paper RWP16-026, Harvard Kennedy School.

Jonas, E., Schimel, J., Greenberg, J., and Pyszczynski, T. (2002). The Scrooge Effect: Evidence That Mortality Salience Increases Prosocial Attitudes and Behavior. *Personality and Social Psychology Bulletin*, 28(10):1342–1353.

Jost, J. T., Langer, M., and Singh, V. (2017). The Politics of Buying, Boycotting, Complaining, and Disputing: An Extension of the Research Program by Jung, Garbarino, Briley, and Wynhausen. *Journal of Consumer Research*, 44(3):503–510.

Kahn, B. E. and Schmittlein, D. C. (1989). Shopping trip behavior: An empirical investigation. *Marketing Letters*, 1(1):55–69.

Kam, C. D. and Deichert, M. (2019). Boycotting, Buycotting, and the Psychology of Political Consumerism. *The Journal of Politics*, pages 000–000.

Khan, R., Misra, K., and Singh, V. (2013). Ideology and Brand Consumption. *Psychological Science*, 24(3):326–333.

Koch, M. T. and Nicholson, S. P. (2016). Death and Turnout: The Human Costs of War and Voter Participation in Democracies. *American Journal of Political Science*, 60(4):932–946.

Kriner, D. L. and Shen, F. X. (2010). *The Casualty Gap: The Causes and Consequences of American Wartime Inequalities*. Oxford University Press, Oxford, England.

Kriner, D. L. and Shen, F. X. (2012). How Citizens Respond to Combat Casualties: The Differential Impact of Local Casualties on Support for the War in Afghanistan. *Public Opinion Quarterly*.

Levay, K. E., Freese, J., and Druckman, J. N. (2016). The Demographic and Political Composition of Mechanical Turk Samples. *SAGE Open*, 6(1):1–17.

Levendusky, M. S. (2017). Americans, Not Partisans: Can Priming American National Identity Reduce Affective Polarization? *The Journal of Politics*, 80(1):59–70.

Liu, J. and Smeesters, D. (2010). Have You Seen the News Today? The Effect of Death-Related Media Contexts on Brand Preferences. *Journal of Marketing Research*, 47(2):251—-262.

Maheswaran, D. (1994). Country of Origin as a Stereotype: Effects of Consumer Expertise and Attribute Strength on Product Evaluations. *Journal of Consumer Research*, 21(2):354–365.

Martin, B. A., Lee, M. S. W., and Lacey, C. (2011). Countering negative country of origin effects using imagery processing. *Journal of Consumer Behaviour*, 10(2):80–92.

Muniz, Albert M., J. and O'Guinn, T. C. (2001). Brand Community. *Journal of Consumer Research*, 27(4):412–432.

Mutz, D. C. (2018). Status threat, not economic hardship, explains the 2016 presidential vote. *Proceedings of the National Academy of Sciences*, 115(19):E4330–E4339.

Nelson, L. J., Moore, D. L., Olivetti, J., and Scott, T. (1997). General and Personal Mortality Salience and Nationalistic Bias. *Personality and Social Psychology Bulletin*, 23(8):884–892.

Samiee, S., Shimp, T. A., and Sharma, S. (2005). Brand origin recognition accuracy: Its antecedents and consumers' cognitive limitations. *Journal of International Business Studies*, 36(4):379–397.

Schildkraut, D. J. (2011). *Americanism in the Twenty-First Century: Public Opinion in the Age of Immigration*. Cambridge University Press, New York.

Schildkraut, D. J. (2014). Boundaries of American Identity: Evolving Understandings of "Us". *Annual Review of Political Science*, 17(1):441–460.

Sedikides, C. (1993). Assessment, enhancement, and verification determinants of the self-evaluation process. *Journal of Personality and Social Psychology*, 65(2):317–338.

Sedikides, C. and Strube, M. J. (1997). Self-evaluation: To thine own self be good, to thine own self be sure, to thine own self be true, and to thine own self be better. In *Advances in Experimental Social Psychology*, volume 29, pages 209–269.

Shachar, R., Erdem, T., Cutright, K. M., and Fitzsimons, G. J. (2011). Brands: The Opiate of the Nonreligious Masses? *Marketing Science*, 30(1):92–110.

Sharma, S., Shimp, T. A., and Shin, J. (1995). Consumer ethnocentrism: A test of antecedents and moderators. *Journal of the Academy of Marketing Science*, 23(1):26–37.

Shimp, T. A. and Sharma, S. (1987). Consumer Ethnocentrism: Construction and Validation of the CETSCALE. *Journal of Marketing Research*, 24(3):280–289.

Siverson, R. M. and Emmons, J. (1991). Birds of a Feather: Democratic Political Systems and Alliance Choices in the Twentieth Century. *Journal of Conflict Resolution*, 35(2):285–306.

Tajfel, H. and Turner, J. C. (1979). An integrative theory of intergroup conflict. In Austin, W. G. and Worchel, S., editors, *The Social Psychology of Intergroup Relations*, pages 33–47. Brooks/Cole Publishing Company, Monterey, CA.

Valentino, N. A., Hutchings, V. L., and Williams, D. (2004). The Impact of Political Advertising on Knowledge, Internet Information Seeking and Candidate Preference. *Journal of Communication*, 54(2):337–354.

A Appendix

A.1 Casualties and Elite Cues

Experimental research emphasizes elite cues as a key driver of public assessment of war. When elite consensus is strong, the public is relatively unaware of war's costs like causalities and defer to politicians (Berinsky, 2009). Though our baseline finding somewhat challenges this argument – Americans consistently responded to local casualties – elite cues may still factor into how casualties strengthen national identification.

We analyze the mediating role of elite cues in local casualty response using weekly campaign advertising during the 2006 midterm election. Iraq War-related ads were elite cues quasi-randomly distributed with respect to local casualties in the same week. The lead time needed to purchase ad time constrained candidates in targeting ads based on recent casualty exposure. The correlation between local cumulative casualty exposure and Iraq-related campaign ads is less than .05 in absolute value and irrespective of local partisanship.

We draw on the Wesleyan Advertising Project's (WAP) advertisement-appearance-level data, which reports each ad spot run by candidates running for the House of Representatives and Senate in 2006 for the top 100 designated market areas (DMAs) in the US. For each spot, the WAP reports its length, air date, DMA, the associated candidate and political party, and a wealth of other information about the ad's content. Political advertising during the 2006 midterms began in January and ran through Election Day in November, so we focus on this time period.[1] Television ads, especially in the context of the 2006 election, are highly salient political messages to which a large majority of the population is exposed (Valentino et al., 2004; Gerber et al., 2011). We focus on ads related to the Iraq War, as coded by the WAP, as they are most likely to contain elite cues about the character and direction of ongoing military conflict. For each DMA-week, we measure the percentage of total political advertising time devoted to ads related to the Iraq War.[2] We then match these DMA-weeks

[1] The campaign season generally may have primed American national identity, but this effect would obtain across large swaths of the country and would not be correlated with casualty exposure.

[2] This could represent a particular candidate running the same ad repeatedly in a given week, or over

to store-weeks based on a DMA-ZIP code crosswalk. This measure allows us to explore if localized elite cues augment, or possibly diminish, the effect of local casualty exposure on national identity attachment.

We also take into account the partisan dynamics of political advertising, subsetting the data by partisan lean using 2000 George W. Bush county-level vote share. We split the data into Safe Democratic areas (in which Bush lost by 10 percentage points or more), Safe Republican areas (in which Bush won by 10 percentage or more), and Tossup areas (in which the vote margin was less than 10 percentage points). This strategy allows us to see how areas with a given ex ante level of partisanship respond differently to cues about the direction of the conflict.[3] We also calculate the percentage of Iraq War-related advertising separately for Democratic and Republican candidates as a further extension, allowing us to see whether consumers respond differently to partisan messaging on the direction of military conflict.

An additional consideration is that the WAP advertising data do not represent a geographic census of campaign advertising in 2006, because it excludes areas that are not in the 100 most populous DMAs. In other words, because the WAP data forces us to scale down our sample only to the 100 most populous DMAs, the sample from which we estimate the effects of elite cues is likely to be less nationally representative than the one we use for our baseline models. This is particularly important when we discuss differential effects by local partisanship. The focus on the top 100 DMAs means that we have ample coverage for Democratic-leaning areas, but comparatively less for tossup areas, and much less coverage for Republican-leaning areas. Our store sample size is cut nearly in half.[4] As a result, while we may be best positioned to understand the effect of elite cues in Democratic-leaning areas, we are more cautious in interpreting our results in Republican-leaning and tossup areas.

We use a similar modeling strategy as we use in our baseline analyses. However, to avoid

multiple weeks.

[3]The total amount of televised political advertising varies predictably by the partisanship of the area in 2006. The average amount of weekly televised political advertising was about 90 minutes in Democratic-leaning areas, 55 minutes in Republican-leaning areas, and 105 minutes in tossup areas.

[4]IRI tracks supermarket scanner data for 320 unique stores in Democratic-leaning areas, 194 unique stores in Republican-leaning areas, and 227 unique stores in tossup areas within the top 100 DMAs.

modeling and interpreting a three-way interaction between casualty exposure, perceived brand nationality, and a time-varying measure of advertising, we subset our data in 2006 to include only product-category store-weeks with above-average $AmericanScore_i$. This means that we keep only observations with $AmericanScore_i > 4$.[5] We then estimate a series of models, in line with our baseline specifications, that include interactions between local casualty exposure and advertising on the Iraq War.

Appendix Table A.12 presents the advertising model results. Each column represents identically specified models for different samples: Column (1) represents all areas, column (2) represents Safe Democratic areas, column (3) represents Safe Republican areas, and column (4) represents tossup areas. These models are specified in line with our baseline cumulative casualty models. The coefficients of interest are those on the interaction between logged cumulative local casualties and the share of advertising time spent discussing Iraq. Generally, the results suggest that greater advertising on Iraq augments the national identification effect that local casualty exposure has. This effect appears to be strongest and most consistent in Safe Democratic areas, but also reaches statistical significance at $p < .1$ in Tossup areas. Meanwhile, increased advertising on Iraq in Safe Republican areas has no detectable interactive effect with cumulative casualty exposure. Again, however, because the sample is restricted to the 100 most populous DMAs, there is sparser coverage of Republican-leaning areas than Democratic-leaning areas. The extent to which more urban Republican areas differ from Republican-leaning areas at large may also limit the generalizability of any inferences on these areas.

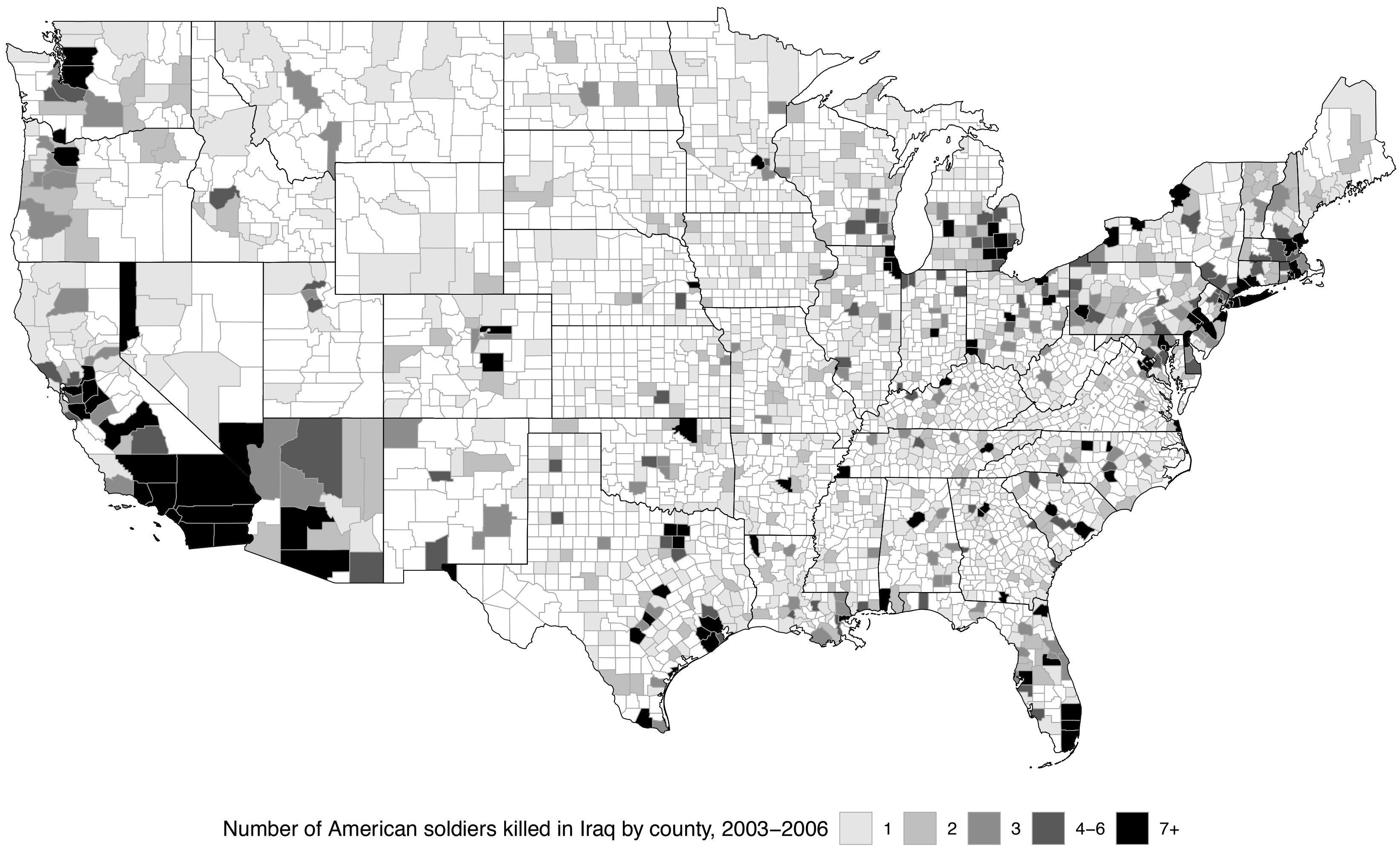
Appendix Table A.12 explores the effect of total advertising time spent discussing Iraq, but this effect could be heterogeneous by the partisan identity of the messenger. Appendix Table A.13 separates out the share of local advertising dedicated to the Iraq War by party. These results suggest that Democratic advertising on the Iraq War is largely driving the interactive relationship between casualty exposure and advertising in Safe Democratic and

[5]In other words, this strategy pools all product-category store-weeks with above-average perceived American origin.

Tossup areas. Meanwhile, our tentative conclusions in Republican-leaning districts are quite different. While relatively more Democratic advertising on Iraq has no effect on change in sales of American-sounding brands in Safe Republican areas, relatively more Republican advertising in those same areas areas appears to diminish the national identity-augmenting effect of local casualty exposure. However, this result is only statistically significant at $p < 0.1$, and this result is potentially limited by the fact that we have non-representative coverage of Republican-leaning areas.

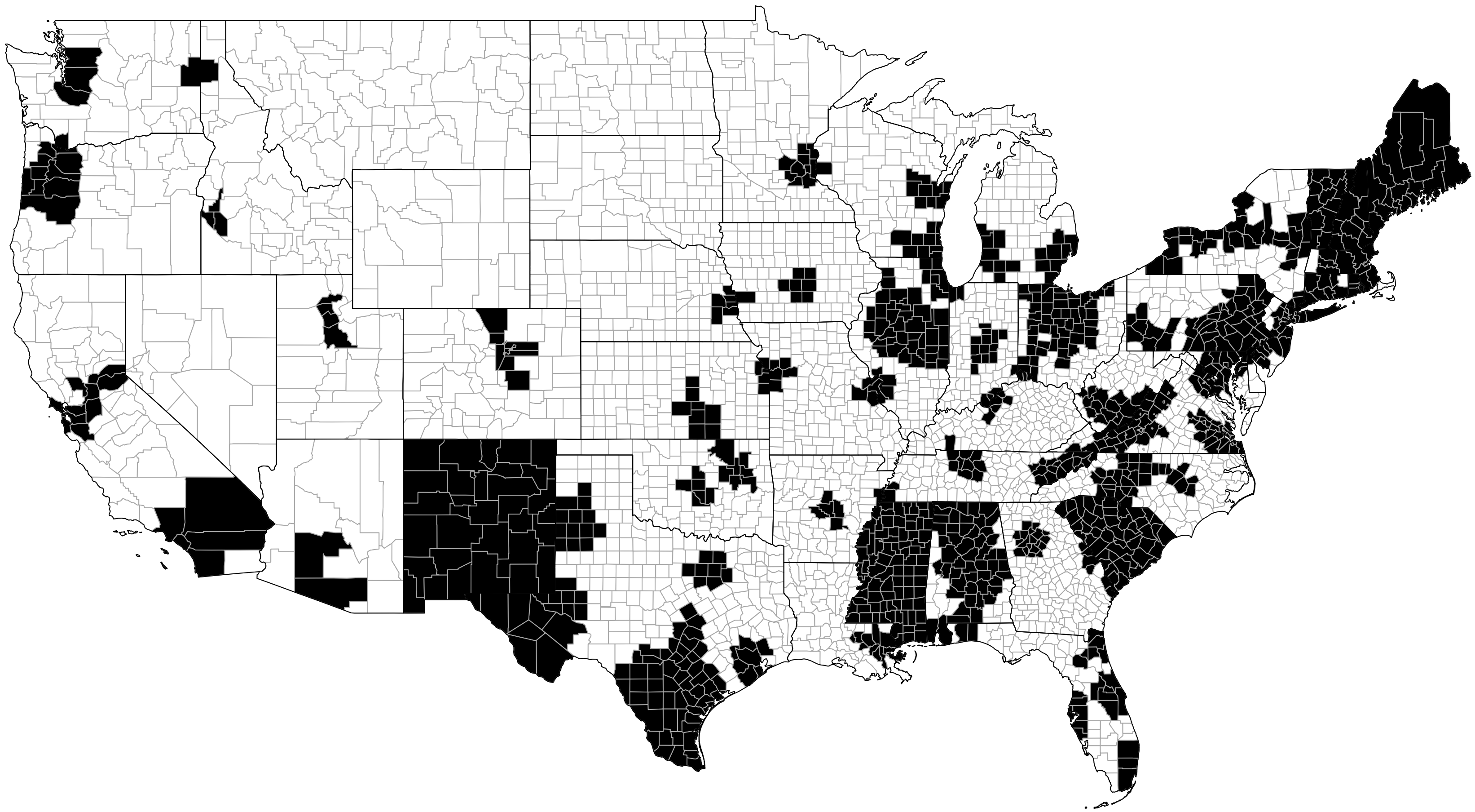
Overall, these results suggest that elite cues play a meaningful role in shaping the salience of local casualty exposure. Our inferences are relatively strongest in Democratic-leaning areas, in which increases in advertising on the Iraq War augmented the national identification-enhancing effect of cumulative local casualty exposure. We find a similar result in Tossup areas, although this inference stands on slightly weaker ground. Meanwhile, the results diverge in Republican-leaning areas, in which Republican advertising on Iraq diminishes the effect that local casualty exposure has on national identification.

Appendix Figure A.1: Distribution of American Soldiers Killed in Iraq by County, 2003-2006



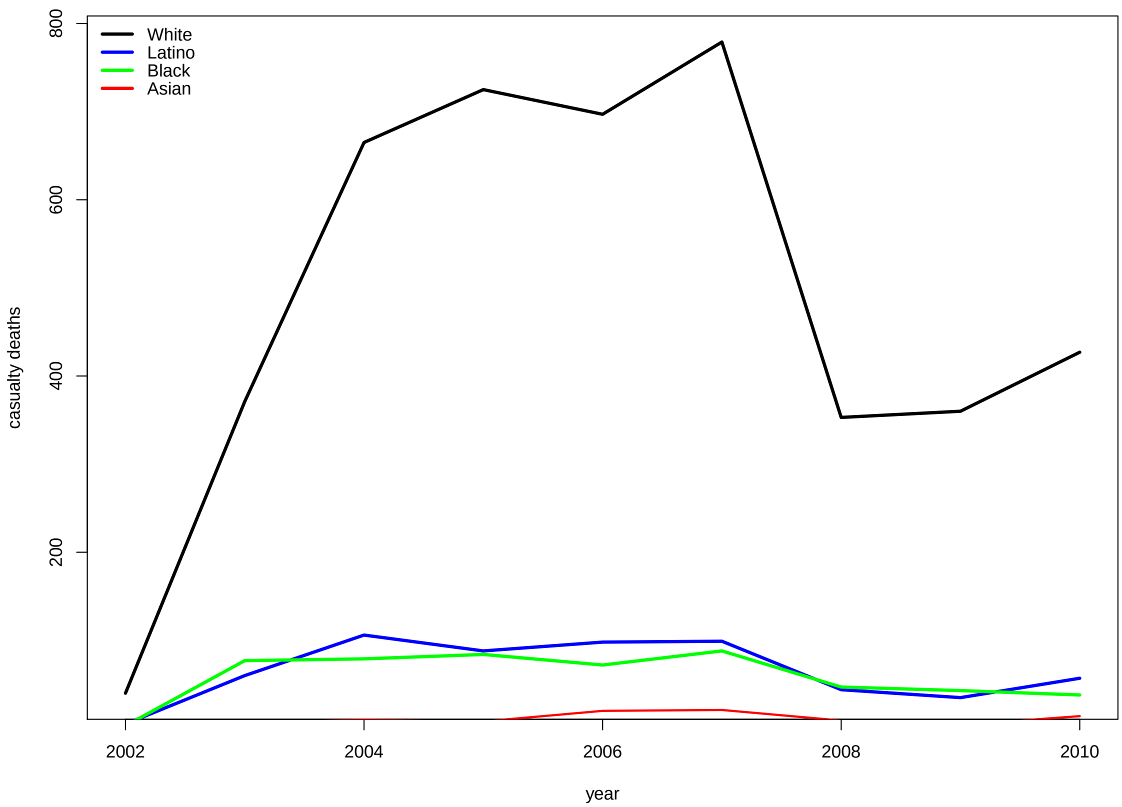
Data Source: Associated Press.

Appendix Figure A.2: Geographic Coverage of IRI Supermarket Scanner Data



Counties in black are in one of 50 geographic markets for which IRI maintains data.

Appendix Figure A.3: Yearly Iraq War Casualties by Race



Data Source: US Department of Defense.

Appendix Table A.1: American Cities with Most Casualties in Iraq, 2003-2006

City	State	Iraq casualties (2003-2006)
New York	New York	29
Houston	Texas	23
Los Angeles	California	22
San Antonio	Texas	22
Phoenix	Arizona	17
Fort Worth	Texas	14
Philadelphia	Pennsylvania	12
Portland	Oregon	12
Las Vegas	Nevada	11
Miami	Florida	11
Tucson	Arizona	11
Austin	Texas	10
Baltimore	Maryland	10
El Paso	Texas	10
Jacksonville	Florida	10
San Diego	California	10
Buffalo	New York	9
Tampa	Florida	9
Cincinnati	Ohio	8
Columbus	Ohio	8
Mesa	Arizona	8
Virginia Beach	Virginia	8

Data Source: Associated Press

Appendix Table A.2: States with Most Casualties in Iraq, 2003-2006

State	Iraq casualties (2003-2006)
California	311
Texas	265
Pennsylvania	146
New York	138
Ohio	131
Florida	128
Michigan	118
Illinois	111
Georgia	89
Virginia	86

Data Source: Associated Press.

Appendix Table A.3: Service Unit Locations with Most Casualties in Iraq, 2003-2006

Unit Location	State	Iraq casualties (2003-2006)
Camp Pendleton	California	304
Fort Hood	Texas	290
Camp Lejeune	North Carolina	242
Fort Campbell	Kentucky	175
Fort Carson	Colorado	124
Fort Bragg	North Carolina	103
Fort Stewart	Georgia	103
Twentynine Palms	California	96
Fort Lewis	Washington	88
Fort Riley	Kansas	69

Data Source: Associated Press.

Appendix Table A.4: Distribution of Perceived American Brands by Product Category

Category	Total Number of Brands	Number of Brands with $AmericanScore_i \geq 3$	Percent of Brands with $AmericanScore_i \geq 3$
Beer	2730	1078	39 %
Blades	129	101	78 %
Carbonated Beverages	512	336	66 %
Cigarettes	779	548	70 %
Coffee	538	179	33 %
Cold Cereal	647	564	87 %
Deodorant	232	131	56 %
Diapers	61	40	66 %
Facial Tissue	68	49	72 %
Frozen Dinners	337	194	58 %
Frozen Pizza	213	83	39 %
Hot Dogs	344	216	63 %
Laundry Detergent	125	94	75 %
Margarine / Butter	86	60	70 %
Mayonnaise	105	49	47 %
Mustard / Ketchup	428	227	53 %
Paper Towel	49	44	90 %
Peanut Butter	67	47	70 %
Razors	36	32	89 %
Salty Snacks	1663	778	47 %
Shampoo	1209	533	44 %
Soup	199	73	37 %
Spaghetti Sauce	410	81	20 %
Sugar Substitute	71	54	76 %
Toilet Tissue	70	44	63 %
Tooth Brush	350	135	39 %
Tooth Paste	215	169	79 %
Yogurt	274	124	45 %

Appendix Table A.5: IRI Market Names

Atlanta, GA
Birmingham/Montgomery, AL
Boston, MA
Buffalo/Rochester, NY
Charlotte, NC
Chicago, IL
Cleveland, OH
Dallas, TX
Des Moines, IA
Detroit, MI
Grand Rapids, MI
Green Bay, WI
Harrisburg/Scranton, PA
Hartford, CT
Houston, TX
Indianapolis, IN
Kansas City, MO
Knoxville, TN
Los Angeles, CA
Milwaukee, WI
Minneapolis/St. Paul, MN
Mississippi
New England
New Orleans, LA
New York, NY
Oklahoma City, OK
Omaha, NE
Peoria/Springfield, IL
Philadelphia, PA
Phoenix, AZ
Portland, OR
Providence, RI
Raleigh/Durham, NC
Richmond/Norfolk, VA
Roanoke, VA
Sacramento, CA
Salt Lake City, UT
San Diego, CA
San Francisco, CA
Seattle/Tacoma, WA
South Carolina
Spokane, WA
St. Louis, MO
Syracuse, NY
Toledo, OH
Tulsa, OK
Washington, DC
West Texas/New Mexico

Appendix Table A.6: Weekly Casualties and American Brand Share, 2003-2006

	$\Delta ShareYear\text{-}2001_{ijkt}$			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$LocalCasualty_{jt} * AmericanScore_i$	0.00014**	0.00012**	0.00011**	0.00030***
	(0.00006)	(0.00005)	(0.00006)	(0.00006)
$NationalCasualties_t * AmericanScore_i$	0.000001	−0.000002**	0.00001***	−0.000001
	(0.000001)	(0.000001)	(0.000001)	(0.000002)
$LocalCasualty_{jt}$	0.00025	−0.00095***	−0.00050*	−0.00152***
	(0.00029)	(0.00023)	(0.00026)	(0.00026)
$NationalCasualties_t$	0.00001	−0.00001	−0.00001**	0.00005***
	(0.000004)	(0.000004)	(0.00001)	(0.00001)
$AmericanScore_i$	−0.00013***	−0.00052***	−0.00122***	−0.00091***
	(0.00003)	(0.00003)	(0.00003)	(0.00004)
$HomePrice_{jt}$	0.00004***	0.00002***	−0.00002***	−0.00003***
	(0.000003)	(0.000003)	(0.000003)	(0.000002)
$HomePrice_{jt} * AmericanScore_i$	−0.00001***	−0.000005***	0.000002***	0.000005***
	(0.000001)	(0.000001)	(0.000001)	(0.000001)
$Enlistment_j$	−0.00105***	−0.00261***	−0.00331***	−0.00348***
	(0.00027)	(0.00029)	(0.00032)	(0.00033)
$Enlistment_j * AmericanScore_i$	−0.00013**	0.00027***	0.00059***	0.00060***
	(0.00006)	(0.00006)	(0.00007)	(0.00007)
$Enlistment_j * AmericanScore_i$	0.0000002***	0.0000003***	0.0000002***	0.0000002***
	(0.00000)	(0.00000)	(0.00000)	(0.0000000)
$\Delta PriceYear\text{-}2001_{ijkt}$	−0.00053***	−0.00052***	−0.00040***	−0.00019***
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
$\Delta Variants2006\text{-}2001_{ijkt}$	0.00900***	0.00857***	0.00872***	0.00968***
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
$Intercept$	−0.00042***	0.00127***	0.00433***	0.00358***
	(0.00012)	(0.00014)	(0.00016)	(0.00019)
Observations	6,715,772	6,344,222	5,756,986	5,533,301
R^2	0.10797	0.12753	0.13501	0.14029

Note: *p<0.1; **p<0.05; ***p<0.01

Appendix Table A.7: Cumulative Casualties and American Brand Share, 2003-2006

	$\Delta ShareYear\text{-}2001_{ijkt}$			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * AmericanScore_i$	0.00021***	0.00018***	0.00012***	0.00010***
	(0.00002)	(0.00001)	(0.00001)	(0.00001)
$ln(CumulNatCasualties_{jt}) * AmericanScore_i$	−0.00004***	−0.00088***	0.00052***	0.00090***
	(0.000005)	(0.00004)	(0.00009)	(0.00015)
$ln(CumulCasualties_{jt})$	0.00006	−0.00097***	−0.00081***	−0.00060***
	(0.00010)	(0.00007)	(0.00007)	(0.00007)
$ln(CumulNatCasualties_{jt})$	0.00004*	0.00338***	0.00286***	0.00724***
	(0.00002)	(0.00019)	(0.00043)	(0.00066)
$AmericanScore_i$	0.00007**	0.00544***	−0.00506***	−0.00809***
	(0.00003)	(0.00028)	(0.00071)	(0.00115)
$HomePrice_{jt}$	0.00004***	0.00003***	−0.00001**	−0.00002***
	(0.000003)	(0.000003)	(0.000003)	(0.000003)
$HomePrice_{jt} * AmericanScore_i$	−0.00001***	−0.00001***	0.0000002	0.000003***
	(0.000001)	(0.000001)	(0.000001)	(0.000001)
$Enlistment_j$	−0.00062**	−0.00139***	−0.00211***	−0.00255***
	(0.00028)	(0.00030)	(0.00034)	(0.00035)
$Enlistment_j * AmericanScore_i$	−0.00026***	0.000004	0.00036***	0.00042***
	(0.00006)	(0.00007)	(0.00007)	(0.00008)
$Population_{j2000}$	0.0000001***	0.0000004***	0.0000003***	0.0000002***
	(0.0000000)	(0.0000000)	(0.0000000)	(0.0000000)
$\Delta PriceYear\text{-}2001_{ijkt}$ (2001-Year)	−0.00053***	−0.00052***	−0.00040***	−0.00019***
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.00901***	0.00857***	0.00872***	0.00968***
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
$Intercept$	−0.00062***	−0.02187***	−0.01712***	−0.05248***
	(0.00015)	(0.00128)	(0.00323)	(0.00524)
Observations	6,715,772	6,344,222	5,756,986	5,533,301
R^2	0.10801	0.12760	0.13511	0.14046
Note:				*p<0.1; **p<0.05; ***p<0.01

Appendix Table A.8: Cumulative Casualties Weighted by Population and American Brand Share, 2003-2006

	$\Delta ShareYear\text{-}2001_{ijkt}$			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$CumulCasualties_{jt}/Pop_{j2000} * AmericanScore_i$	0.01682***	0.00577**	0.00838***	0.00483***
	(0.00607)	(0.00277)	(0.00225)	(0.00178)
$ln(CumulNatCasualties_{jt}) * AmericanScore_i$	−0.00003***	−0.00079***	0.00055***	0.00095***
	(0.000005)	(0.00004)	(0.00009)	(0.00015)
$CumulCasualties_{jt}/Population_{j2000}$	0.00723	−0.02747**	−0.09463***	−0.07052***
	(0.02824)	(0.01273)	(0.01033)	(0.00816)
$ln(CumulNatCasualties_{jt})$	0.00004**	0.00286***	0.00275***	0.00726***
	(0.00002)	(0.00019)	(0.00043)	(0.00066)
$AmericanScore_i$	−0.00001	0.00486***	−0.00532***	−0.00846***
	(0.00003)	(0.00028)	(0.00071)	(0.00115)
$HomePrice_{jt}$	0.00005***	0.00002***	−0.00001***	−0.00002***
	(0.000003)	(0.000003)	(0.000002)	(0.000002)
$HomePrice_{jt} * AmericanScore_i$	−0.00001***	−0.000004***	0.000002***	0.00001***
	(0.000001)	(0.000001)	(0.000001)	(0.000001)
$Enlistment_j$	−0.00020	−0.00176***	−0.00291***	−0.00309***
	(0.00027)	(0.00028)	(0.00031)	(0.00032)
$Enlistment_j * AmericanScore_i$	−0.00011*	0.00029***	0.00060***	0.00064***
	(0.00006)	(0.00006)	(0.00007)	(0.00007)
$\Delta PriceYear\text{-}2001_{ijkt}$	−0.00054***	−0.00053***	−0.00040***	−0.00019***
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.00901***	0.00857***	0.00872***	0.00968***
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
$Intercept$	−0.00071***	−0.01847***	−0.01615***	−0.05256***
	(0.00015)	(0.00127)	(0.00323)	(0.00524)
Observations	6,715,772	6,344,222	5,756,986	5,533,301
R^2	0.10792	0.12751	0.13509	0.14045
Note:				*p<0.1; **p<0.05; ***p<0.01

Appendix Table A.9: Cumulative Casualties and Coalition of the Willing Brand Share, 2003-2006

	$\Delta ShareYear\text{-}2001_{ijkt}$			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * CoalitionScore_i$	-0.00025^{***}	-0.00016^{***}	-0.00005^{**}	-0.00008^{***}
	(0.00003)	(0.00002)	(0.00002)	(0.00002)
$ln(CumulNatCasualties_{jt}) * CoalitionScore_i$	0.00003^{***}	0.00040^{***}	-0.00045^{***}	-0.00041^{*}
	(0.00001)	(0.00007)	(0.00015)	(0.00025)
$ln(CumulCasualties_{jt})$	0.00084^{***}	-0.00111^{***}	-0.00141^{***}	-0.00022^{***}
	(0.00010)	(0.00008)	(0.00008)	(0.00008)
$ln(CumulNatCasualties_{jt})$	-0.00012^{***}	0.00305^{***}	0.01411^{***}	0.01139^{***}
	(0.00002)	(0.00019)	(0.00045)	(0.00072)
$CoalitionScore_i$	0.00007	-0.00223^{***}	0.00410^{***}	0.00418^{**}
	(0.00005)	(0.00045)	(0.00114)	(0.00195)
$HomePrice_{jt}$	-0.00003^{***}	-0.00003^{***}	-0.000003	-0.00001^{***}
	(0.000003)	(0.000003)	(0.000003)	(0.000003)
$HomePrice_{jt} * CoalitionScore_i$	-0.000002^{*}	-0.000003^{***}	-0.00001^{***}	-0.00001^{***}
	(0.000001)	(0.000001)	(0.000001)	(0.000001)
$Enlistment_j$	-0.00345^{***}	-0.00071^{**}	-0.00005	0.00114^{***}
	(0.00028)	(0.00032)	(0.00036)	(0.00038)
$Enlistment_j * CoalitionScore_i$	0.00008	-0.00023^{**}	-0.00037^{***}	-0.00061^{***}
	(0.00009)	(0.00011)	(0.00012)	(0.00013)
$Population_{j2000}$	0.000001^{***}	0.000001^{***}	0.000001^{***}	0.000001^{***}
	(0.0000000)	(0.0000000)	(0.0000000)	(0.0000000)
$\Delta PriceYear\text{-}2001_{ijkt}$	-0.00086^{***}	-0.00061^{***}	-0.00056^{***}	-0.00027^{***}
	(0.00002)	(0.00001)	(0.00001)	(0.00001)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.00767^{***}	0.00744^{***}	0.00729^{***}	0.00681^{***}
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
$Intercept$	-0.00100^{***}	-0.02290^{***}	-0.10558^{***}	-0.09102^{***}
	(0.00015)	(0.00133)	(0.00336)	(0.00572)
Observations	6,573,689	5,968,261	5,472,115	5,056,532
R^2	0.09648	0.10082	0.10790	0.10488

Note: *p<0.1; **p<0.05; ***p<0.01

Appendix Table A.10: Cumulative Casualties and France/Germany Brand Share, 2003-2006

	Δ*ShareYear-2001*$_{ijkt}$			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * France + GermScore_i$	−0.00024***	−0.00008***	−0.00007***	−0.00005*
	(0.00003)	(0.00002)	(0.00002)	(0.00003)
$ln(CumulNatCasualties_{jt}) * France + GermScore_i$	0.00005***	0.00039***	0.00115***	0.00040
	(0.00001)	(0.00007)	(0.00016)	(0.00026)
$ln(CumulCasualties_{jt})$	0.00056***	−0.00118***	−0.00119***	−0.00039***
	(0.00011)	(0.00009)	(0.00009)	(0.00009)
$ln(CumulNatCasualties_{jt})$	−0.00021***	0.00340***	0.00990***	0.01145***
	(0.00002)	(0.00021)	(0.00048)	(0.00077)
$France + GermScore_i$	−0.00095***	−0.00340***	−0.00938***	−0.00385*
	(0.00006)	(0.00048)	(0.00123)	(0.00206)
$HomePrice_{jt}$	−0.00002***	−0.00002***	0.0000002	−0.000004
	(0.000003)	(0.000003)	(0.000003)	(0.000003)
$HomePrice_{jt} * France + GermanyScore_i$	0.000004***	0.000004***	0.000003**	0.000002*
	(0.000001)	(0.000001)	(0.000001)	(0.000001)
$Enlistment_j$	−0.00246***	−0.00139***	−0.00047	−0.00002
	(0.00031)	(0.00034)	(0.00038)	(0.00040)
$Enlistment_j * France + GermanyScore_i$	−0.00042***	−0.00037***	−0.00029**	−0.00028**
	(0.00010)	(0.00012)	(0.00013)	(0.00014)
$Population_{j2000}$	0.000001***	0.000001***	0.000001***	0.000001***
	(0.0000000)	(0.0000000)	(0.0000000)	(0.0000000)
Δ*PriceYear-2001*$_{ijkt}$	−0.00116***	−0.00095***	−0.00093***	−0.00069***
	(0.00002)	(0.00002)	(0.00002)	(0.00002)
Δ*VariantsYear-2001*$_{ijkt}$	0.00781***	0.00762***	0.00757***	0.00732***
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
Intercept	0.00071***	−0.02280***	−0.07151***	−0.08826***
	(0.00016)	(0.00143)	(0.00362)	(0.00605)
Observations	4,770,144	4,339,951	3,966,649	3,662,239
R^2	0.08881	0.09496	0.10040	0.09930
Note:				*p<0.1; **p<0.05; ***p<0.01

Appendix Table A.11: Cumulative Casualties and Demographic Variation - 2006

	$\Delta Share2006\text{-}2001_{ijkt}$								
	Native-born	Blue collar	White collar	Black	Asian	Armed forces	Income	Age 18-39	Hispanic
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$ln(CumulCasualties_{jt}) * AmerScore_i * NativeBorn_j$	0.00016***								
	(0.00001)								
$ln(CumulCasualties_{jt}) * AmerScore_i * NativeBorn_j^2$	−0.000001***								
	(0.00000)								
$ln(CumulCasualties_{jt}) * AmerScore_i * BlueCollar_j$		0.00001***							
		(0.00000)							
$ln(CumulCasualties_{jt}) * AmerScore_i * WhiteCollar_j$			−0.000008***						
			(0.00000)						
$ln(CumulCasualties_{jt}) * AmerScore_i * Black_j$				−0.000004***					
				(0.000001)					
$ln(CumulCasualties_{jt}) * AmerScore_i * Asian_j$					−0.000006***				
					(0.000002)				
$ln(CumulCasualties_{jt}) * AmerScore_i * ArmedForces_j$						0.000008			
						(0.000009)			
$ln(CumulCasualties_{jt}) * AmerScore_i * HHIncome_j$							0.000000		
							(0.000000)		
$ln(CumulCasualties_{jt}) * AmerScore_i * Age18\text{-}39_j$								−0.00021	
								(0.00013)	
$ln(CumulCasualties_{jt}) * AmerScore_i * Hispanic_j$									0.000002
									(0.000002)
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	5,030,086	5,030,086	5,030,086	5,030,086	5,030,086	5,030,086	5,030,086	5,030,086	5,030,086
R^2	0.143	0.143	0.143	0.143	0.143	0.143	0.143	0.143	0.143

Note: $^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$

Appendix Table A.12: Casualty Exposure and Political Advertising - 2006

	$\Delta Share2006\text{-}2001_{ijkt}$			
	All	Safe Dem.	Safe Rep.	Tossup
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * IraqAds_{jt}$	0.00001***	0.00002***	−0.00001	0.00001*
	(0.00000)	(0.00000)	(0.00001)	(0.00001)
$ln(CumulCasualties_{jt})$	0.001***	0.0004***	0.001***	0.001***
	(0.00005)	(0.0001)	(0.0001)	(0.0001)
$IraqAds_{jt}$	−0.00001**	−0.00003***	0.00004**	−0.00001
	(0.00001)	(0.00001)	(0.00001)	(0.00001)
$TotalAdvertising_{jt}$	−0.000	0.00000***	−0.00000***	−0.00000***
	(0.000)	(0.000)	(0.000)	(0.000)
$HomePrice_{jt}$	0.00000	0.00002***	−0.00001**	−0.00004***
	(0.00000)	(0.00000)	(0.00001)	(0.00000)
$Enlistment_{j}$	0.0001	−0.002***	0.001**	0.002***
	(0.0002)	(0.0003)	(0.001)	(0.001)
$\Delta Price2006\text{-}2001_{ijkt}$ (2001-2006)	−0.0002***	0.00004*	−0.001***	−0.0004***
	(0.00002)	(0.00002)	(0.00004)	(0.00003)
$\Delta Variants2006\text{-}2001_{ijkt}$	0.011***	0.012***	0.011***	0.011***
	(0.00002)	(0.00003)	(0.00004)	(0.00003)
$Intercept$	−0.002***	−0.003***	−0.002***	−0.002***
	(0.0001)	(0.0002)	(0.0002)	(0.0002)
Observations	2,401,268	1,042,164	623,928	735,176
R^2	0.128	0.125	0.132	0.130
Note:				*p<0.1; **p<0.05; ***p<0.01

Appendix Table A.13: Casualty Exposure and Partisan Political Advertising - 2006

	$\Delta Share2006\text{-}2001_{ijkt}$			
	All	Safe Dem.	Safe Rep.	Tossup
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * DemIraqAds_{jt}$	0.00001***	0.00002***	−0.00000	0.00001**
	(0.00000)	(0.00000)	(0.00001)	(0.00001)
$ln(CumulCasualties_{jt}) * RepIraqAds_{jt}$	−0.00002*	−0.00001	−0.00003*	−0.00002
	(0.00001)	(0.00002)	(0.00002)	(0.00004)
$ln(CumulCasualties_{jt})$	−0.00002***	−0.00003***	0.00001	−0.00001
	(0.00001)	(0.00001)	(0.00002)	(0.00001)
$DemIraqAds_{jt}$	0.0001***	0.00003	0.0001***	0.00004
	(0.00002)	(0.00004)	(0.00003)	(0.0001)
$RepIraqAds_{jt}$	−0.000	0.00000***	−0.00000***	−0.00000***
	(0.000)	(0.000)	(0.000)	(0.000)
$TotalAdvertising_{jt}$	0.00000	0.00002***	−0.00001**	−0.00004***
	(0.00000)	(0.00000)	(0.00001)	(0.00000)
$HomePrice_{jt}$	0.0001	−0.002***	0.001*	0.002***
	(0.0002)	(0.0003)	(0.001)	(0.001)
$Enlistment_{j}$	−0.0002***	0.00004*	−0.001***	−0.0004***
	(0.00002)	(0.00002)	(0.00004)	(0.00003)
$\Delta Price2006\text{-}2001_{ijkt}$	0.011***	0.012***	0.011***	0.011***
	(0.00002)	(0.00003)	(0.00004)	(0.00003)
$\Delta Variants2006\text{-}2001_{ijkt}$	0.001***	0.0005***	0.001***	0.001***
	(0.00005)	(0.0001)	(0.0001)	(0.0001)
$Intercept$	−0.002***	−0.003***	−0.002***	−0.002***
	(0.0001)	(0.0002)	(0.0002)	(0.0002)
Observations	2,401,268	1,042,164	623,928	735,176
R^2	0.128	0.125	0.132	0.130
Note:				*p<0.1; **p<0.05; ***p<0.01