War on Aisle 5: Casualties, National Identity, and Consumer Behavior*

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Abstract

How does national identity shape domestic reactions to international politics? Although prominent debates in International Relations (IR) turn on this question, scholars lack empirical approaches to establish the casual effects of national identity on behavior. We introduce an approach that leverages strong similarities between political and consumer behavior. Sales of American supermarket brands is a time-varying measure of national identity's effects on choice behavior. We find that when American soldiers died in Iraq, the market share of American brands grew in their US hometowns. Casualties produce feelings of threat. Share growth was stronger in supermarkets frequented by Republicans, who are more prone to embrace national identity in response to threat. Partisan cues and animosity towards erstwhile allies do not drive our findings. Our study strengthens IR's theoretical microfoundations by establishing, with high external validity, that national identity shapes mass responses to international politics.

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1 Introduction

How does national identity – the sense of belonging to one's nation – shape how people react to international relations? Prominent puzzles in International Relations (IR) scholarship highlight this question. For example, trade competition fuels support for far-right parties, but mechanisms are unclear (Margalit 2019; Naoi 2020). Analyses of voting behavior assume a preference for protectionist policies (Colantone and Stanig 2018), but trade policy attitudes reflect factors like ethnonationalism, a belief in the superiority of one's own nation (Mansfield and Mutz 2013). Likewise, national security issues can influence vote choice through voters' cost-benefit analysis (Karol and Miguel 2007), and by reinforcing national identity (Huddy et al. 2005; Koch and Nicholson 2016). These puzzles reflect deeper questions about how material self-interest and social identity shape mass reactions to international politics (Mutz 2018; Herrmann 2017).

These puzzles persist because we are unable to directly test if international politics shapes domestic politics *because* it strengthens national identity. Existing strategies leave a gap between expressed sentiment and observable behavior. Scholars use survey experiments to analyze causal drivers of foreign policy sentiment such as identity, emotion, and information (Hafner-Burton et al. 2017; Kertzer and Tingley 2018). Sentiment, however, may diverge from behavior if surveys prompt reflection on otherwise low-salience issues or provide information that would be otherwise unavailable (Barabas and Jerit 2010; Guisinger 2017). Analyses of electoral outcomes leverage quasi-random exposure to phenomena like import competition or war casualties (Kriner and Shen 2009; Autor et al. 2020). While such analyses establish the causal causal effects of international politics on formal political behavior, they are not are designed to parse mechanisms.

We introduce a novel empirical strategy to evaluate the role of national identity in mass responses to international politics. Our strategy relies on a well-documented pattern in American consumer behavior: Americans express their national identity by purchasing American brands (Shimp and Sharma 1987). These are brands that consumers perceive as American by virtue of name or symbolism such as Sam Adams (beer), Cracker Jack (salty snacks), and American Heritage (condiments). We propose that, if international politics strengthens national identity, the market share of American brands should grow. An observable, causally-identified change in American brand purchases requires sufficient awareness of international politics and an activation of national identity strong enough to influence how people spend their own money. Evidence of mass consumer response provides greater confidence in analogous electoral responses to international politics.

Consumer behavior is an insightful proxy for political behavior. Similar to vote choice, brand choice reflects a combination of material and non-material preferences shaped by price, quality, early socialization, peer influence, and advertising (Bronnenberg and Dubé 2017); and psychological processes that operate outside conscious awareness (Maheswaran 1994; Gürhan-Canli and Maheswaran 2000). Social identity-based branding makes brand choice a tangible expression of consumers' identity (White and Dahl 2006, 2007). We focus on supermarket purchases, which unlike most political behaviors, is a frequent, consistent, well-documented, and nearly universal behavior in the US (Kahn and Schmittlein 1989; Sorensen et al. 2017). Supermarket scanner data report weekly sales and other brand- and store-level information for representative samples of stores. As compared to other approximations of real-world political behavior, such as survey and laboratory experiments, analysis of consumer behavior provides strong external validity of identity-based mechanisms.

We deploy this strategy to analyze the effect of local Iraq War casualties on the local weekly market share of American brands during 2003-2006. Public response to causalities is a core IR research question that speaks to the foundations of democratic accountability during war (Gartner 2008; Kriner and Shen 2010). For a given store-week, a local casualty is the death in that week of an American soldier whose hometown is located in the same county as the store. We hypothesize that local casualties increase local market share of American brands. Casualties activate national identity because they represent a threat to the values and ideals for which soldiers gave their lives (Althaus and Coe 2011) and prime thoughts of death (Koch and Nicholson 2016). Consumers switch to American brands to reaffirm their national identity, a psychological response to buffer against feelings of threat.

We analyze weekly supermarket sales for a nationally representative sample of more than 8,000 brands in over 1,100 US supermarkets. Our original measure of brands' perceived American nationality approximates how consumers infer a brand's country of origin (Samiee et al. 2005). Our identifying assumption is that casualties are quasi-randomly distributed across counties, conditional on military enlistment. For each store-week in 2003-2006, we estimate the change in market share of American brands relative to the same store-week in 2001, the first year for which scanner data are available. This year-over-2001 design provides a pre-Iraq War baseline to evaluate change and holds constant slow-moving drivers of brands' market share in a given store, including customer demographics, ex ante demand for American brands, and product characteristics. Local casualties influence consumption only by activating national identity; casualties did not affect supply, product characteristics, or systematically produce calls to purchase American brands. We control for store-week change in price, retailers' only real-time response to demand shocks, and availability. We separately analyze customer demographic data to evaluate sources of heterogeneity in responses to local casualties.

We find that in store-weeks exposed to casualties, the market share of American brands consistently grew relative to the same week in 2001 on average. The finding is robust to a variety of measurement and model specifications, and controls for store, week, and local characteristics. Cumulative casualty exposure produced even larger market share growth on average. Analyzing customer demographic data, we show that stores with a higher proportion of customers in skilled occupations exhibit a weaker response to local causalities. This finding is consistent with less ethnocentrism among more educated people. We also find a stronger response in stores located in more Republican-leaning counties.

We further unpack these mechanisms. Casualty response exhibits in-group bias rather than out-group animosity. Market shares of brands associated with France and Germany, who opposed the war, declined as much as those linked to the "Coalition of the Willing" allied countries. We further probe the effects of partisanship by leveraging the quasi-random coincidence of local casualties and local Iraq War-related campaign advertising – a type of partisan cue – during the 2006 Congressional elections. Ads alone had no effect but marginally increased share growth in stores with higher cumulative casualties.

Our empirical approach establishes a tighter link between sentiment and behavior, which strengthens the microfoundations of "second image-reversed" IR theories (Gourevitch 1978). Our findings more precisely demonstrate national identity as a mechanism through which international relations influences mass behavior. Canonical theories of foreign policy and democratic accountability, such as theories of interstate crisis bargaining (Fearon 1994) and trade protectionism (Grossman and Helpman 1994), hinge on how and why citizens react to international events. Mass reactions may be consistent with the rational/materialist assumptions of many prevailing IR theories, or reflect psychological processes such as identity that survey research brings to the fore. The relative importance of these mechanisms can influence the robustness of democratic accountability.

Analysis of consumption complements existing empirical approaches. It addresses external validity and replicability concerns about experiments (Hafner-Burton et al. 2017) while also directly testing identity-related mechanisms. In the conclusion, we discuss how our findings bolster the importance of non-material drivers of foreign policy preferences; demonstrate mass awareness of foreign policy in real-world settings; and suggest inconsistencies between sentiment and behavior.

We recognize that consumer behavior and high-frequency consumer data are unfamiliar to our discipline so we highlight key features of our empirical strategy. Our approach is valid for estimating real-time effects, and is analogous to priming in experiments. We show consumption change in the same week as the casualty and, for robustness, lagged by one week. We cannot precisely estimate longer lags because unobserved, non-price responses to demand shocks are more likely. For example, retailers may target casualty-exposed areas with advertisements that cue national identity, a response that requires time to implement. Our results establish the lower and upper bounds of persistence. Our baseline analysis assumes that casualties' effects dissipate completely the following week. Our analysis of logged cumulative casualties assumes that effects persist at the same strength for all subsequent weeks. Given our four-year sample period, this range suggests that activation of national identity could persist for long enough to influence voters' electoral choices.

Effect sizes, the scale of which are unfamiliar in our discipline, are substantively large in the context of year-over-2001 store market share growth. To provide an intuitive baseline, we interpret this growth relative to a one-standard-deviation price drop, the most consistent correlate of market share growth (Yang et al. 2003).[1] Some may be curious about aggregate financial impact of casualty-induced changes in consumption behavior. Such effects are, in our view, secondary to the goal of establishing mass behavioral responses to international politics, and are difficult to precisely estimate in our setting. We regard our interpretations with respect to price changes as a more insightful quantity of interest with respect to micro-level political behavior.

2 International Politics and Consumer Behavior

In our theoretical framework, social identity connects international politics to consumer behavior. People define themselves by their most important social identities (Tajfel and Turner 1979; Tajfel 1981) and derive self-esteem and security from adhering to associated norms and behaviors (Terry et al. 1999; Akerlof and Kranton 2000). National identity is a particularly salient dimension of social identity. Members of national in-groups adhere to prescriptive national identity norms and feel more positive emotions after conforming to prescribed behaviors (Huddy and Khatib 2007).

[1] Consumer price indices rely on similar estimates from weekly scanner data (Hawkes and Piotrowski 2003).

The purchase of domestic brands is a prescriptive behavior associated with national identity. Brand choices generally reflect consumers' most salient social identities (Belk 1988; Shachar et al. 2011). Brands cue nationality with names and symbols that consumers associate with their countries (Verlegh and Steenkamp 1999). Much like social stereotypes, nationality branding cues operate largely outside of consumers' conscious awareness (Martin et al. 2011). Consumers connect domestic brands with moral obligations to advance the economic well-being of co-nationals (Shimp and Sharma 1987; Sharma et al. 1995).

People further embrace their most relevant social identities in response to threats to their safety and self-esteem (Branscombe and et al. 1999; Davies et al. 2008). Reinforcing important social identities reaffirms one's value and purpose (Greenberg and et al. 1992; Arndt and et al. 1997). As Huddy et al. (2005, p. 594) note, "[o]ne of the most pervasive and powerful effects of threat is to increase intolerance, prejudice, ethnocentrism, and xenophobia, regardless of whether threat is defined as a widely acknowledged external force or a subjective, perceived state." In the domain of international politics, events like terrorism and war can raise concerns about threats to personal safety and national security (Huddy et al. 2002; Hetherington and Suhay 2011). Threats to self-esteem emerge with economic dislocation from globalization (Ballard-Rosa et al. 2020, 2021) and broader geopolitical shifts (Mutz 2018).

We argue that sales of domestic brands increase following international political events that are perceived as threatening. Specific consumption-related mechanisms include emotional reactions to threat that increase reliance on heuristics like brand nationality (Maheswaran and Chen 2006) and undermine the security and stability that underpins habitual consumption (Rindfleisch et al. 2009).[2] Our empirical analysis focuses on local Iraq War casualties. The American public public linked the war to the threat of international terrorism (Gershkoff and Kushner 2005). Casualties also generated vivid images that tied death to American identity, such as flag-draped coffins returning to US soil (Gartner 2011). Re-

[2]Consumer anamosity towards brands from foreign adversaries exhibits similar dynamics (Klein 2002).

minders of death induce stronger in-group bias (Greenberg et al. 1997). Koch and Nicholson (2016) argue that war casualties drive voter turnout for this reason. In experiments, death-related cues increase subjects' preference for domestic brands of grocery products such as beverages and candy (Nelson and et al. 1997; Friese and Hofmann 2008; Liu and Smeesters 2010).

We expect that Americans' propensity to switch to American brands varies based on demographic characteristics and other dimensions of social identity. These factors correlate with the propensity to perceive threats, respond with stronger national identity, and change consumption behavior. Empirical tests of heterogeneity in consumption response are stringent because characteristics that influence the propensity to switch also influence baseline preferences. For example, Americans with a stronger sense of national identity have a stronger baseline preference for American brands (Shimp and Sharma 1987). Thus, we would not observe a change in their brand choice even if they more inclined to strengthened national identity in response to threat.

Nevertheless, we note demographic and identity-based factors that may color American's response to international threats. Ideology and partisan identity correspond to distinctive personality traits, cognitive processes, motives, and values, all of which shape how people both perceive international politics and their willingness to change their consumption (Jost 2017). Experimental evidence shows that more conservative Americans respond to threats by switching to brands heavy in national symbolism (Cutright et al. 2011; Shepherd et al. 2015). Conservatives, however, are ex ante more likely to consume American brands (Carney et al. 2008) and have stickier consumption preferences (Khan et al. 2013). We examine two additional traits that can reflect stronger ex ante attachments to American identity: percent of store customers born in the US and the percent employed by the military.

Education has significant implications for both responses to threat and consumption patterns more generally. More educated individuals are less likely to draw rigid in-group/out-group boundaries, (e.g. Sharma et al. 1995; Hainmueller and Hiscox 2010) and are less likely to reaffirm national identity in threatening circumstances (Schildkraut 2014). We expect that more educated Americans are less likely to switch to American brands following local casualties.

Expectations about racial heterogeneity are less clear. Non-white Americans are less likely to reaffirm national identity in response to threat (Schildkraut 2014). Current research finds no racial disparities in the effect of casualties on war support (Gartner and Segura 2000)[3]

3 Local Exposure to Iraq War Casualties, 2003-2006

The Iraq War generated 3,240 US military casualties during 2003-2006, about 90 percent of total US war casualties during these years. Figure 1 plots weekly Iraq casualty counts. The weekly average is relatively stable, with the largest spikes corresponding to predictable moments like the initial invasion and the 2004 insurgency.

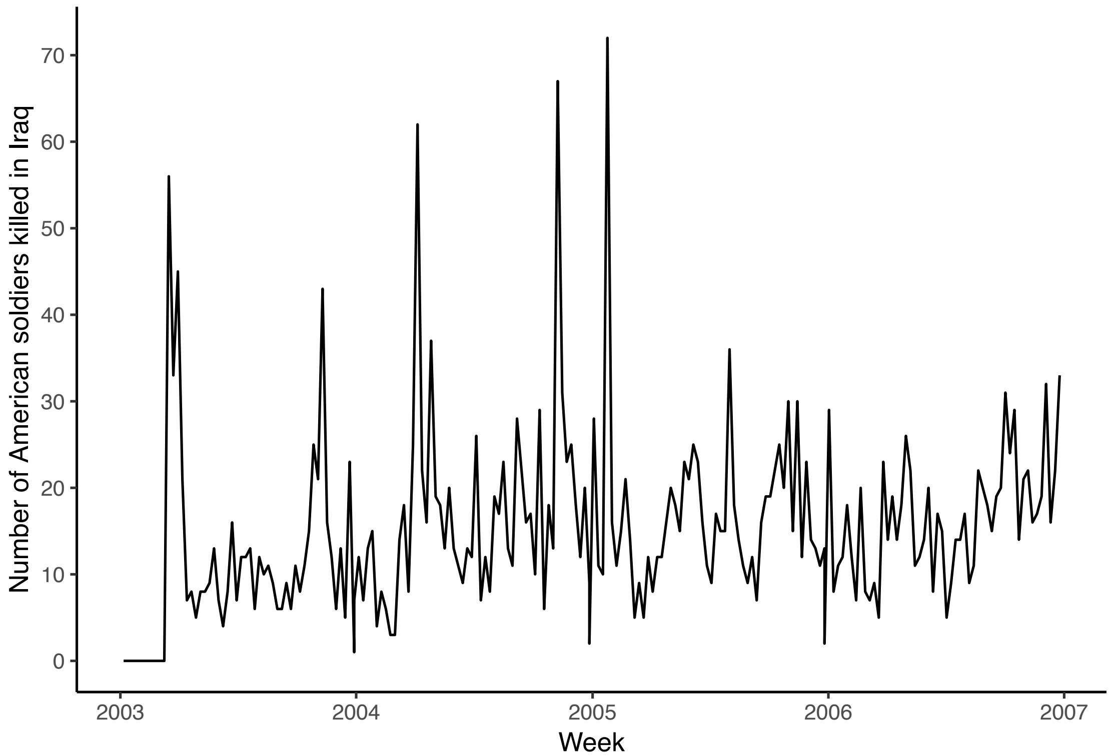
From the perspective of a given US county, a "local" casualty refers to the death of a deployed soldier originally from that county. Approximately 2,000 unique American cities experienced at least one war casualty during the sample period; 18 cities had ten or more casualties.[4] Figure 2 maps cumulative county-level American causalities in Iraq during 2003-2006. The geographic distribution of casualties is roughly proportional to population.

Our identifying assumption is that conditional on military enlistment, a county's exposure to war casualties is quasi-random. Timing of local casualties is clearly exogenous to soldiers' hometowns. Military enlistment, however, is non-random, because the US had an all-volunteer military during the Iraq War. We control for enlistment in all analyses. Americans demonstrated high awareness of local Iraq War casualties. Proximity increases the likelihood of exposure to information about the casualty. Local casualties produce the

[3] Iraq War casualties were overwhelmingly white (data available on request), so we do not anticipate varying responses based on soldiers' race.

[4] See Appendix Tables A.1 (p. A4), A.2 (p. A4), and A.3 (p. A5) for additional details on the distribution of casualties across cities, states, and service units.

Figure 1: Weekly Iraq War Casualties, 2003-2006



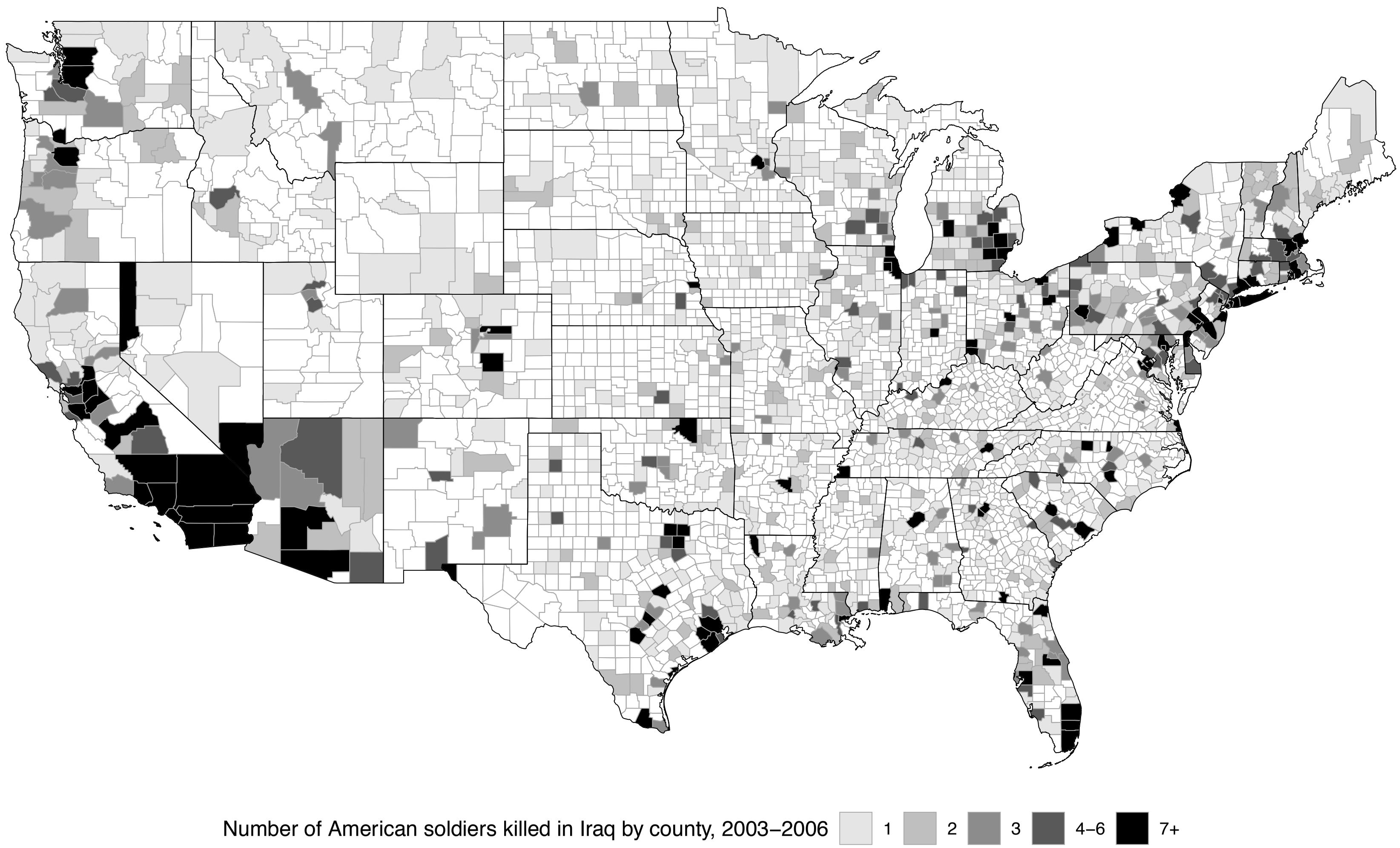
Data Source: Associated Press.

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 implausibly high fraction of survey respondents who report knowing an Iraq War casualty suggests that the mass public perceives strong personal connection to casualties (Gartner 2009).

4 Data and Measurement

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

Figure 2: Distribution of American Soldiers Killed in Iraq by County, 2003-2006



Data Source: Associated Press.

4.1 Weekly Supermarket Purchases

We measure consumer response to US casualties in Iraq using weekly supermarket scanner data supplied by Information Resources Inc. (IRI), a leading source of scanner data (Bronnenberg et al. 2008).[5] These data cover a representative sample of 1,145 supermarkets across 50 IRI-designated geographic markets.[6] Appendix Figure A.2 (p. A3) 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.[7] The 968 counties in the IRI data account for 90 percent of US Iraq War casualties during 2003-2006.

We construct our store-level measures of consumer response using weekly unit sales for 8,644 brands across 27 categories of supermarket products.[8] Major supermarket chains stock mature brands and maintain a relatively stable portfolio of brands within each store. 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 entry or attrition. We aggregate data across multiple stock keeping unit (SKU) codes of a single brand-product category, but not across distinct, but related, brands.[9]

4.2 Brand Nationality

Consumers systematically infer brand nationality from marketing cues rather than searching for official country-of-origin labels (Samiee et al. 2005; Balabanis and Diamantopoulos 2011). For American consumers, brands that incorporate American geographic locations or

[5] These are academic-use data whose use is subject to a confidentiality agreement. The agreement forbids disclosure of brand- and store-specific information.

[6] IRI set its market definitions in 1987 to achieve a representative sample of US consumers, making it unlikely that our findings are an artefact of sample selection.

[7] In 2003 supermarkets accounted for roughly 70 percent of all US grocery purchases (Market Share Reporter, 2003).

[8] 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.

[9] 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.

cultural symbols imply that product is made in the US. By contrast, brand names based on foreign languages provide cues like distinctive letter combinations and special characters (like umlauts and accent marks) that do not occur in English. Some brand names may not incorporate nationality cues but consumers associate them with America because of advertising cues or because the brand itself has become a symbol of national identity. "Objective" brand origin is not relevant for our purposes but inspection of US trademark filings for our sample brands does suggest that brands perceived as American are often trademarks owned by US-based companies.

To measure a product's perceived nationality, we administered surveys via Amazon's Mechanical Turk (MTurk).[10] 1,203 participants received a randomly selected brand name and its product category. We asked "What nationality does this brand most make you think of?" and offered ten possible responses (American, Chinese, English, French, German, Italian, Japanese, Spanish, "none," and "other.")[11] We paid participants per evaluation and restricted each participant to 20 evaluations. Each of the 8,644 brands in our data had seven independent evaluations.

We use these data to construct an index of each brand's perceived American origin. For each independent evaluation, we create an indicator variable, $American_{rb}$, that equals one if respondent r coded brand b as American and zero otherwise. For each brand, we then sum $American_{rb}$ for all seven independent evaluations into a single additive index. Our measure of perceived American nationality is the count of respondents who perceived brand b as American. $AmericanScore_b$ ranges in value from 0 (i.e. no respondents coded brand b as American) to 7 (i.e. all respondents coded brand b as American).

[10] MTurk results are often more reliable results than convenience- and lab-based samples (Berinsky et al. 2012; Huff and Tingley 2015). While MTurk samples may not mirror American demographics (Levay et al. 2016), the gap is likely uncorrelated with brand nationality perceptions.

[11] Brand names included all special characters like accents and umlauts. We also 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.

Table 1: Brand Examples Across AmericanScore$_b$ Values

$AmericanScore_b$ Value	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)

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

Table 1 presents examples of brands at each value of $AmericanScore_b$. Brands with $AmericanScore_b$=7 exhibit strong cues of American nationality, including references to US geography and historical figures. Low-scoring brands often have distinctive non-American elements, including non-English words and foreign geographic references. In a separate survey of 400 US-based respondents, we verify that $AmericanScore_b$ positively and strongly correlates with a widely-used psychometric scale of brands' perceived American origin.[12] One potential concern is that gradated increases in $AmericanScore_b$ (i.e. moving from 3 to 4) may not reflect consistent and meaningful rises in perceived American nationality. Our findings are robust to excluding middle categories ($1 < AmericanScore_b < 6$) and focusing on the most clearly American and non-American brands.

[12]See Appendix Section A.1 (p. A15) for further details.

4.3 Exposure to US Iraq War Casualties

Casualty data are from US Department of Defense press releases compiled by the Associated Press. For each casualty, our data include name, hometown, rank, unit, and date and cause of death.[13] We match each casualty to the county in which their hometown is located. The county level is the most conservative measure of exposure we can accurately construct and is standard in existing research (e.g. Kriner and Shen 2009).

$LocalCasualty_{jt}$ is equal to one if at least one casualty occurred in the same county as store j in week t, and zero otherwise. Of the store-weeks that experienced casualties, 92 percent experienced a single casualty.[14] We verify that our results are robust to continuous measures of local weekly casualty exposure. By the end of the sample period, more than 95 percent of sample stores experienced at least one casualty. We also construct $NationalCasualties_t$, the national total of US Iraq War casualties in week t, which captures time-varying characteristics of the war that could influence consumption including events in Iraq, national-level elite rhetoric, and media coverage.

4.4 Additional Control Variables

Our identifying assumption of quasi-random casualty exposure is conditional on military enlistment. We control for total enlistment over the previous five years in store j's ZIP code. These data are based on enlistees' home addresses and cover all military branches.[15] 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.

We also control for weekly change in local economic conditions as proxied by average

[13] We verify that hometown is distinct from the location of the soldier's service unit. For example, among the 294 casualties of soldiers based at Fort Hood, Texas, there are 259 unique hometowns across 51 US states and territories. See Appendix Table A.3 (p. A5).

[14] An additional 7 percent of store-weeks in which a casualty occurred represented two casualties. The maximum number of casualties in a store-week was four, which represent less than 0.5 percent of store-weeks with at least one casualty.

[15] Data are from Kriner and Shen (2010).

home prices in store j's ZIP code and week t.[16] Though casualty exposure is unlikely to correlate with weekly local economic conditions, economic conditions could influence casualty response by changing emotional states and the information environment.[17]

Partisanship may influence the propensity to respond to threats by changing consumption. As a robustness check we control for George W. Bush's 2000 county vote share. Data come from Dave Leip's Electoral Atlas. Finally, we control for logged county population in 2000 in case population size influences casualty information or response. Data are from the 2000 US Census.

5 Empirical Analysis

For each $AmericanScore_i$-category-store-week in our sample, we model the average change in market share 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: 1,145 supermarkets,

k: 27 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-week. Measuring market share, as opposed to the total number of units sold, scales store's sales of a brand relative to overall demand for that product category in that store-week. Changes in market share capture shifts in demand for brands distinct from changes in demand for an

[16]These data are from zillow.com

[17]We verify that American brands are not systematically cheaper such that tighter budget constraints prompt a switch to lower-cost brands.

entire product category. For each 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 from 2003-2006, for every $AmericanScore_i$-category-store-week in our sample, we calculate the change in market share between week t in that year and the same week in 2001 ($ShareYear\text{-}2001_{ijkt}$). Measuring change in demand within each store holds constant all relatively time-invariant baseline characteristics of a store's customer base that influence sales, including *ex ante* customer preferences. Likewise, our specification controls for product category characteristics such as the propensity to have nationality-based branding. The year-over-2001 change accounts for seasonal fluctuations, such as higher demand for American brands around patriotic holidays. We choose 2001, the first year for which scanner data are available, as a baseline because it precedes all Iraq War casualties. If 9/11 and/or the eleven Afghanistan War casualties in 2001 increased sales of American brands, this would bias against our expected finding for subsequent years. We verify that our results hold if we restrict our sample to January-August.

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

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

where

$Year$	$\in$	[2003, 2004, 2005, 2006],
$\Delta ShareYear\text{-}2001_{ijkt}$	=	average difference in market share from 2001 to *Year* for brands with $AmericanScore_i$ in store j belonging to product category k in week t,
$LocalCasualty_{jt}$	=	indicator for US war casualty in Iraq from same county as store j in week t,
$AmericanScore_i$	=	index from 0-7 indicating level of perceived American nationality,
$Enlistment_j$	=	total military enlistment in same ZIP code as store j in last five years,
$HomePrice_{jt}$	=	average home price in same ZIP code as store j and week t,
$Population_{j2000}$	=	population in same county as store j in year 2000,
$\Delta PriceYear\text{-}2001_{ijkt}$	=	average difference in price from 2001 to *Year* for $AmericanScore_i$ in store j belonging to product category k in week t,
$\Delta VariantsYear\text{-}2001_{ijkt}$	=	average difference in number of variants from 2001 to *Year* for $AmericanScore_i$ in store j belonging to product category k in week t, and
ϵ_{ijkt}	=	normally distributed random error term.

The coefficient of interest is β_3, on the interaction between local casualty exposure and $AmericanScore_i$. We interact $Enlistment_j$ and $HomePrice_jt$ with $AmericanScore_i$ to allow for the salience of ex ante military enlistment and weekly fluctuation in local economic conditions to vary based on level of perceived American brand nationality. We control for two time-varying brand-store characteristics that affect market share (Ataman et al. 2010). $\Delta PriceYear\text{-}2001_{ijkt}$ controls for average price changes and the effect of promotional, limited-time discounts.[18] Price promotions are retailers' fastest response to negative demand

[18]We verify that average weekly price changes are uncorrelated with $AmericanScore_i$.

Table 2: Weekly Casualties and American Brand Share - 2003

	Δ *Share2003-2001*$_{ijkt}$ (%)	
	(1)	(2)
$LocalCasualty_{jt} * AmericanScore_i$	0.016***	0.01437**
	(0.00613)	(0.00623)
$NationalCasualties_t * AmericanScore_i$		0.00014
		(0.00009)
$LocalCasualty_{jt}$	0.03231	0.02518
	(0.02815)	(0.02860)
$NationalCasualties_t$		0.00052
		(0.00041)
$AmericanScore_i$	−0.01196***	−0.01350***
	(0.00245)	(0.00265)
Observations	6,715,772	6,715,772
Controls	✓	✓

Note: *p<0.1; **p<0.05; ***p<0.01. All models estimated with OLS. See Appendix Table A.5 (p. A6) for controls.

shocks.[19] Non-price responses, such as advertising, are less likely because they require longer lead times to implement. Retailers' contracts with manufacturers forbid changes to shelf space allocation and location.[20] We also control for average weekly changes in the number of brand SKUs that a store stocks in a product category, $\Delta\,VariantsYear\text{-}2001_{ijkt}$. All else equal, consumers are more likely to purchase a brand if a store stocks more varieties.

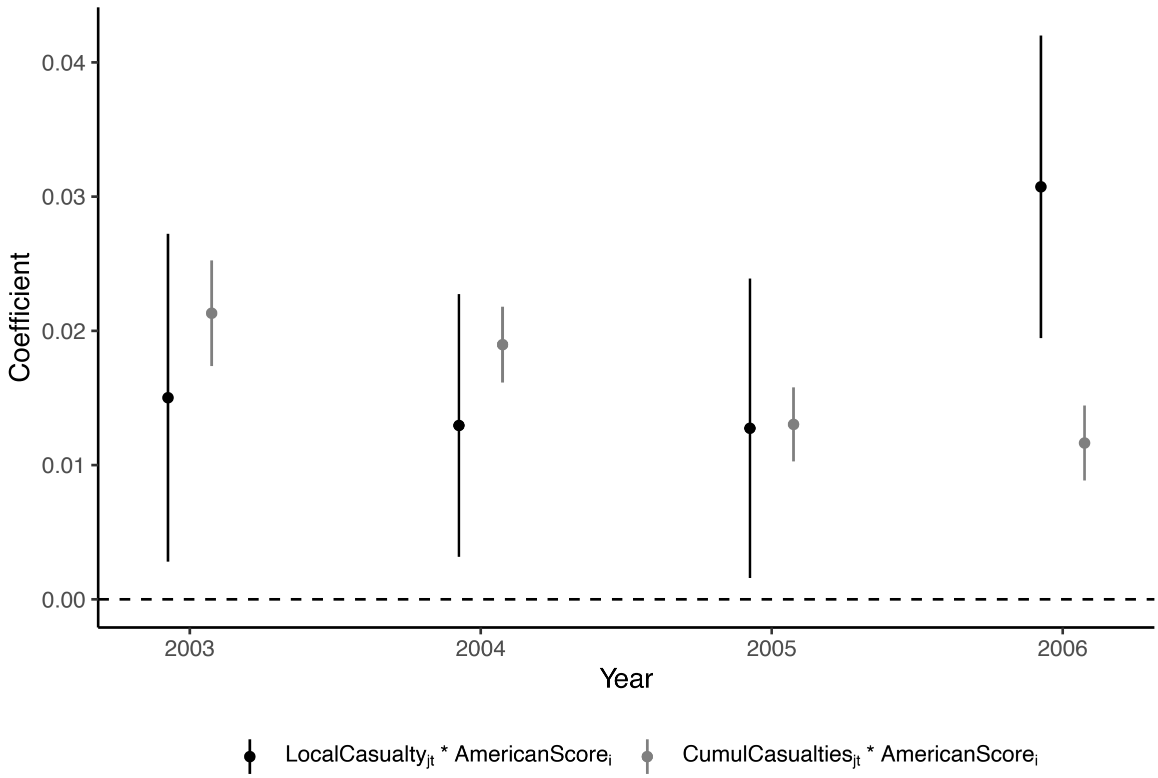
5.1 Baseline Results

Table 2 presents results for 2003, the first year of the Iraq War. Model (1) focuses on local weekly casualties.[21] The coefficient of interest is on $LocalCasualty_{jt} * AmericanScore_i$. This interaction is positive and statistically significant in both models, indicating that, on average, in weeks that stores experience a local casualty, the market share growth of American brands increases. Model (2) adds controls for the national casualty environment, the total number of American casualties in Iraq in week t. Our baseline finding is unchanged.

Figure 3 plots in black the coefficient on $LocalCasualty_{jt} * AmericanScore_i$ for Model (2), estimated annually during 2003-2006. The figure illustrates that Americans consistently responded to local casualties by switching to American brands during 2003-2006. The magnitude of this effect is substantively large. Among store-weeks exposed to casualties, the average $AmericanScore_i = 7$ brand saw market share growth equivalent to a one-standard deviation price drop.

We verify our findings hold across multiple robustness tests.[22] We lag casualty exposure by one week to allow for delay in consumer response (Appendix Table A.6 (p. A7)). We measure casualty exposure from both the Iraq and Afghanistan wars. Separately, we find no effect of just Afghanistan War casualties, which generated 10 percent of causalities during the sample. We control for George W. Bush's 2000 vote share in store j's. Finally, we cluster standard errors by store.

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



Effect of weekly and cumulative casualty exposure on market share of American brands over time, with 95 percent confidence intervals. Coefficients for weekly casualties drawn from Appendix Table A.5 (p. A6). Coefficients for cumulative casualties drawn from Appendix Table A.7 (p. A8).

5.2 Cumulative Casualties

We thus far have assumed that Americans respond to each week's casualties *de novo*. We next evaluate the effect of cumulative casualties. Information-based theories of casualty and public opinion emphasis variation in public recall of past events. We modify our baseline analysis, replacing weekly casualties with $ln(CumulCasualties_{jt})$, the natural log of cumulative local

[19] Manufacturers provide retailers with a trade allowance to finance promotions so price fluctuations are less likely correlated with local characteristics.

[20] 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.

[21] See Appendix Table A.5 (p. A6) for the results for all years, as well as coefficient estimates on control variables.

[22] Results not in the appendix are available upon request.

Table 3: Cumulative Casualties and American Brand Share - 2006

	$\Delta Share2006\text{-}2001_{ijkt}$ (%)	
	(1)	(2)
$ln(CumulCasualties_{jt}) * AmericanScore_i$	0.011***	0.010***
	(0.001)	(0.001)
$ln(CumulNatCasualties_{jt}) * AmericanScore_i$		0.090***
		(0.015)
$ln(CumulCasualties_{jt})$	−0.044***	−0.060***
	(0.007)	(0.007)
$ln(CumulNatCasualties_t)$		0.724***
		(0.066)
$AmericanScore_i$	−0.099***	−0.809***
	(0.003)	(0.115)
Observations	5,533,301	5,533,301
Controls	✓	✓

Note: *p<0.1; **p<0.05; ***p<0.01. All models estimated with OLS. See Appendix Table A.7 (p. A8) for controls.

casualties in the same county as store j from the beginning of the Iraq War to week t. We also control for $ln(CumulNatCasualties_t)$ in line with our baseline specifications. These measures assume a high level of recall and provide an upper bound for the size of consumer response.

Table 3 presents our results for cumulative casualties for 2006, the last year in our sample.[23] 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 increases the market share growth of American brands. 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

[23]See Appendix Table A.7 (p. A8) for the results for all years, as well as coefficient estimates on control variables.

independent of the war's overall trajectory and total costs, and changes in elite rhetoric and access to information. In stores with a one-standard-deviation higher exposure to logged cumulative casualties, average growth in the market share of $AmericanScore_i = 7$ brands is about two-thirds the market share growth of a one-standard-deviation price drop.

Figure 3 displays in gray the coefficient on the interaction between $ln(CumulCasualties_{jt})$ and $AmericanScore_i$ and its 95 percent confidence interval for every year during 2003-2006, extracted from our Model (2) specification. 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.

We perform the same robustness tests as in the baseline analysis: using cumulative Iraq and Afghanistan casualties, cumulative Afghanistan casualties only, controlling for Bush 2000 vote share, and clustering standard errors by store.[24] We also omit counties that were two standard deviations from mean logged cumulative casualties. Finally, our results do not substantively change if we weight local cumulative casualties by county population in 2000 because more populous counties are likely to have higher cumulative casualties.[25]

Finally, we ensure that both our baseline and cumulative casualty results are robust to limiting our sample to the brands that most respondents judged to be American or non-American. We dichotomize $AmericanScore_i$ to create $American_i$, which is equal to 1 if $AmericanScore_i > 5$ and 0 if $AmericanScore_i < 2$. We then run our same baseline and cumulative casualty models on this subsample. Our main results hold (Appendix Tables A.9 (p. A10) and A.10 (p. A11)). This exercise illustrates that our findings are not driven primarily by small or gradated increases in the middle of $AmericanScore_i$, but rather by switching from brands that are clearly non-American to brands that are clearly American.

[24] All results not in the appendix are available upon request.
[25] See Appendix Table A.8 (p. A9).

6 Mechanisms

We hypothesize that local casualties increase sales of American brands as an affirmation of national identity in response to threat. Accordingly, stores whose customers are more likely to affirm national identity should exhibit larger effects. We evaluate these implications using customer demographic data. For each store, we have population characteristics in 2000 for a two-mile radius, the standard catchment area of a supermarket.[26] We have data ob the percent of customers in 2000 employed in skilled jobs ($SkilledOcc_{j2000}$),[27] employed by the military (including civilians) ($ArmedForces_{j2000}$), born in the US ($NativeBorn_{j2000}$), and who are black ($Black_{j2000}$). We supplement these data with county-level 2000 Bush vote share ($Bush_{j2000}$) as a measure of partisanship.

We first establish baseline demographic correlates of American brand market share. These correlations help us interpret change in response to local casualties. We estimate the average market share of products at each level of $AmericanScore_i$ in 2001, the baseline year for our main results. As this is a cross-sectional analysis, we include category, store, and week fixed effects; omit the weeks after September 11, 2001; and exclude counties with populations below and above two standard deviations of the mean. We estimate an OLS model that includes the interaction of $AmericanScore_i$ and each of the customer characteristics.

Appendix Table A.11 (p. A12) summarizes our results. All else equal, American brands have higher market shares in stores with larger proportions of Republican-leaning customers. For $AmericanScore_i = 7$ brands, a five-percentage point increase in 2000 Bush vote share corresponds to a roughly 0.013-percentage-point higher average market share. Stores with high proportions of white collar and armed forces workers also exhibit higher American brand market share. Stores with more US-born customers show no correlation with American brand market share. Stores with higher proportions of black customers exhibit lower market shares.

We next analyze demographic variation in the propensity to change consumption af-

[26]IRI derived demographic data from the 2000 US Census.
[27]These are "white collar" workers as defined in the 2000 US Occupation and Employment Census.

Table 4: Cumulative Casualties and Demographic Variation - 2006

	$\Delta Share2006\text{-}2001_{ijkt}$
	(1)
$ln(CumulCasualties_{jt}) * AmericanScore_i * Bush_{j2000}$	0.00166^{***}
	(0.00015)
$ln(CumulCasualties_{jt}) * AmericanScore_i * SkilledOcc_{j2000}$	-0.00075^{***}
	(0.00019)
$ln(CumulCasualties_{jt}) * AmericanScore_i * ArmedForces_{j2000}$	0.00082
	(0.00089)
$ln(CumulCasualties_{jt}) * AmericanScore_i * NativeBorn_{j2000}$	−0.00009
	(0.00018)
$ln(CumulCasualties_{jt}) * AmericanScore_i * Black_{j2000}$	−0.00010
	(0.00013)
Observations	5,030,086
Controls	✓

Note: *p<0.1; **p<0.05; ***p<0.01. Model estimated with OLS. All constituent interaction terms included in model and suppressed in table. Stores in counties with populations two standard deviations above or below the mean excluded.

ter local casualties. We modify our analysis of cumulative casualties in 2006 by adding three-way interactions between cumulative casualty exposure, American Score, and each demographic variable. Table 4 presents the results. Partisanship is the strongest correlate of American brand share growth. For stores with mean logged cumulative casualty exposure, a five-percentage-point increase in 2000 Bush vote share is associated with a roughly 0.1-percentage-point increase in $AmericanScore_i = 7$ brand market share growth on average. Combined with our baseline analysis, these results show that customers in more Republican areas had a higher ex ante propensity to purchase American brands that local casualties further magnified.[28] Stores with high proportions of skilled worker customers exhibit a

[28] We note that the county partisanship measure is more coarse than the store-level demographic characteristics.

slight decline in market share growth. For stores with mean cumulative casualty exposure, a five-percentage-point increase in skilled occupation customers is associated with a roughly 0.04-percentage-point decrease in market share growth for $AmericanScore_i = 7$ on average. This decline is consistent with education corresponding to weaker propensity to strengthen national identity in response to threat. We find no propensity to switch to American brands in stores with higher proportions of customers born in the US, employed in the armed forces, or who are black. The latter two traits correspond to correlations in baseline demand for these brands, which suggests that ex ante brand choices left less scope for change.

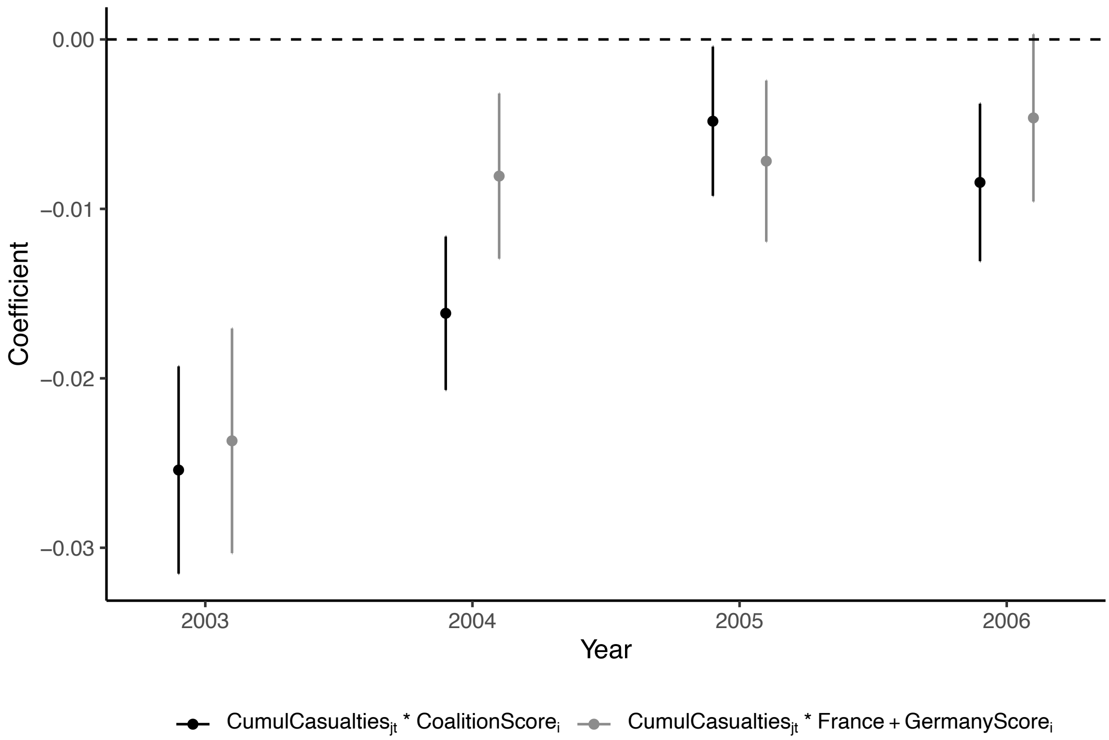
6.1 In-Group Solidarity or Out-Group Animosity?

Debates over IR's microfoundations rest, in part, on ascertaining the variants of national identity that emerge in response to international politics. "Constructive" forms of national identity are consistent with democratic accountability whereas unquestioning blind patriotism likely subverts accountability (Huddy and Khatib 2007). We engage this distinction by analyzing whether strengthened national identity reflected greater anamosity towards foreigners. The Iraq War divided US allies in Europe. France and Germany opposed the invasion while the United Kingdom, Italy, and Spain joined the "Coalition of the Willing by sending troops to Iraq."[29] We estimate change in the market share of brands associated with these two sets of countries. If wartime alliances/division shape affinities/animosity, we would see divergence market share growth between brands perceived to be from coalition countries versus France and Germany. Comparison of these two groups also allows us to differentiate between a proactive shift towards American brands versus a shift incidental to the boycott of French and German brands.

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

[29] Unfortunately, we lack brands associated with Iraq or the Middle East more generally, which is likely an artefact of the product categories in our data.

Figure 4: Casualties Reduce Market Share of Both 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.12 (p. A13). Coefficients for French and German brands drawn from Appendix Table A.13 (p. A14).

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.[30] 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 casualties did not sharpen antipathy towards out-groups even

[30] See Appendix Tables A.12 (p. A13) and A.13 (p. A14) for full results.

as they strengthened national identification.

6.2 Alternate Mechanism: Partisan Cues

We further unpack the role of partisanship by evaluating the effect of partisan cues on casualty response.[31] Elite cues may shape behavior in a range of ways. They may simply be reminders of ongoing war or prime feelings of national identity. They might also shape how ordinary people perceive and process threats to their national identity. Any of these mechanisms are consistent with elite cues coloring localized responses to casualties.

We measure cues using Iraq War-related campaign advertising in the 2006 midterm elections.[32] For a given week, ads were quasi-randomly distributed with respect to local casualties. The lead time needed to purchase television ads precluded real-time targeting based on casualty exposure. 2006 House and Senate advertising data are from the Wesleyan Advertising Project (WAP), which reports each instance of a given candidate-advertisement for the top 100 US media markets, known as designated market areas (DMAs). All ads are 30 seconds. The WAP reports air dates, media market, candidate and political party, and detailed ad content. Political advertising during the 2006 midterms began in January and ran through Election Day in November, so we focus on this time period.[33] 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 and et al. 2011). Ads related to the Iraq War are most likely to contain partisan cues about the character and direction of military conflict. For each media market-week, we measure the percentage of televised campaign ads that the WAP coded as related to the Iraq War. Using a media market-ZIP code crosswalk, we measure store-week exposure to Iraq War-related ads.

An important limitation of this analysis is that the top 100 media markets skew Democrat

[31] We lack data to evaluate the role of social networks, another possible driver of casualty response (Kertzer and Zeitzoff 2017).

[32] Partisan cues may have been weaker in 2006 due to the collapse of partisan consensus about the war (Berinsky 2009). This biases against finding an effect.

[33] 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.

Table 5: Casualty Exposure and Political Advertising - 2006

	Change in market share			
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * IraqAds_{jt} * AmericanScore_i$	0.0002* (0.0001)			
$ln(CumulCasualties_{jt}) * DemIraqAds_{jt} * AmericanScore_i$		0.0003** (0.0001)		0.0003*** (0.0001)
$ln(CumulCasualties_{jt}) * RepIraqAds_{jt} * AmericanScore_i$			−0.001 (0.0004)	−0.001 (0.0004)
$ln(CumulCasualties_{jt}) * AmericanScore_i$	0.013*** (0.001)	0.013*** (0.001)	0.014*** (0.001)	0.013*** (0.001)
Observations	5,326,090	5,326,090	5,326,090	5,326,090
Controls	✓	✓	✓	✓

Note: *p<0.1; **p<0.05; ***p<0.01. All constituent interactions included in model and suppressed in table.

so our sample is not nationally representative with respect to partisanship. Of our sample stores located in the top 100 media markets, 320 are in Democratic-leaning counties, where Bush lost by 10 percentage points or more in 2000; 227 stores are in tossup counties, where Bush won by less than 10 percentage points; and only 194 stores are in Republican-leaning counties, where Bush won by more than ten percentage points. As a result, we are more circumspect in our inferences about consumer response in tossup and Republican-leaning areas.[34] The correlation between local cumulative casualty exposure and Iraq War-related ads is less than .05 in absolute value and irrespective of local partisanship.

We estimate the joint effect of casualties and partisan Iraq War-related cues. We add to our cumulative casualty model a three-way interaction between local cumulative casualty exposure, perceived brand nationality, and our time-varying measure of exposure to to Iraq-

[34]Average total political advertising also correlates with local partisanship. Democratic-leaning areas received 90 minutes, Republican areas 55 minutes, and tossup areas 105 minutes.

related political advertising. We also disaggregate our measure of exposure to advertising by political party. As in the cumulative casualty analysis, we include 2000 county-level Bush vote share and its interaction with $AmericanScore_i$ to control for local partisanship.

Table 5 presents the advertising model results. Taken together, the results indicate that partisan cues modestly magnify casualties' effects on brand choice but do not independently activate national identity. Regardless, across every regression, the coefficient on the interaction between cumulative casualties and $AmericanScore_i$ continues to be positive and statistically significant on its own. The magnifying effects of advertising are strongest in Democrat-leaning areas though data limitations preclude stronger inferences about heterogeneity in the effect of partisan cues. Overall, these results suggest that elite cues play a meaningful, but not necessarily paramount, role in shaping the salience of local casualty exposure.

7 Conclusion

Our findings demonstrate that international politics can drive mass behavior by strengthening national identity. Consumption behavior is an insightful proxy for political behavior, especially regarding the causal effects of social identity on real-world behavior. Brand choice is a tangible and readily observable manifestation of consumers' social identity. We have focused on national identity but in the U.S., product branding incorporates many politically-relevant identities including gender, race/ethnicity, and geographic location. The weekly frequency of supermarket scanner data can measure virtually real-time activation of social identities following exogenous shocks. The universality of consumption captures responses from those excluded or alienated from formal political participation. Taken together, analysis of mass consumption behavior is a powerful complement to experimental research on sentiment and analysis of electoral outcomes.

This study also bolsters the importance of culture/identity-related mechanisms in IR by showing that non-material preferences drive material decisions. Whereas experiments typically yield cross-sectional evidence, our research design based on high-frequency data reveals when national identity becomes relatively more important than other dimensions of preferences. We contribute to a growing body of research that moves away from stark dichotomies between material and non-material preferences towards more nuanced accounts that map their interaction (Ballard-Rosa et al. 2020, 2021; **?**). Our novel measure of the revealed strength of national identity complements existing measures including psychometric scales, discourse/text analysis, and ethnography (Huddy and Khatib 2007; Abdelal et al. 2009).

Further, we contribute to debates about information and foreign policy attitudes by showing that in the real-world contexts that people receive and interpret information about international politics, they are sufficiently moved to change their behavior. Our research design captures the sum of information effects, including information characteristics – source, content, frequency – and the numerous factors that influence receptivity, interpretation, and the propensity to change behavior. Local casualties consistently changed behavior during 2003-2006 despite fluctuations in partisan cues and media coverage about the Iraq War (Baum and Potter 2008; Berinsky 2009).

More generally, our findings suggest that sentiment and behavior reflect different cognitive processes. Local casualties consistently activated national identity during 2003-2006 even as public support for the Iraq War declined.[35] We conjecture that our findings reflect processes that operate outside conscious awareness. Strengthened national identity manifests in a domain that relies heavily on such processes and is ostensibly unrelated to war. Consumer behavior provides an empirical tool that can help clarify the role of identity-based processes in mass response to external threat (Levendusky 2017; Myrick 2021). More broadly, this divergence underscores the importance of both sentiment and behavior to fully specifying the microfoundations of IR theory.

[35] Appendix Figure A.1 (p. A2) summarizes shifts in public support. Casualties may influence public support for war via casualty aversion (Gartner 2008), strengthened resolve (Gelpi et al. 2009), social identity (Althaus and Coe 2011) and information.

References

Abdelal, Rawi, Yoshiko M. Herrera, Alastair Iain Johnston, and Rose McDermott, *Measuring Identity: A Guide for Social Scientists*, Cambridge University Press, April 2009.

Akerlof, George A. and Rachel E. Kranton, "Economics and Identity," *The Quarterly Journal of Economics*, August 2000, *115* (3), 715–753.

Althaus, Scott L. and Kevin Coe, "Social Identity Processes and the Dynamics of Public Support for War," *Public Opinion Quarterly*, January 2011, *75* (1), 65–88.

_ , Brittany H. Bramlett, and James G. Gimpel, "When War Hits Home: The Geography of Military Losses and Support for War in Time and Space," *Journal of Conflict Resolution*, 2012, *56* (3), 382–412.

Arndt, Jamie and et al., "Suppression, Accessibility of Death-Related Thoughts, and Cultural Worldview Defense: Exploring the Psychodynamics of Terror Management," *Journal of Personality and Social Psychology*, 1997, *73* (1), 5–18.

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

Autor, David, David Dorn, Gordon Hanson, and Kaveh Majlesi, "Importing Political Polarization? The Electoral Consequences of Rising Trade Exposure," *American Economic Review*, 2020, *110* (10), 3139–3183.

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

Ballard-Rosa, Cameron, Amalie Jensen, and Kenneth Scheve, "Economic Decline, Social Identity, and Authoritarian Values in the United States," *International Studies Quarterly*, 2021. Forthcoming.

_ , Mashail Malik, Stephanie Rickard, and Kenneth Scheve, "The Economic Origins of Authoritarian Values: Evidence from Local Trade Shocks in the United Kingdom," *Comparative Political Studies*, 2020, *Forthcoming.*

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

Baum, Matthew A. and Philip B.K. Potter, "The Relationships Between Mass Media, Public Opinion, and Foreign Policy: Toward a Theoretical Synthesis," *Annual Review of Political Science*, 2008, *11* (1), 39–65.

Belk, Russell W., "Posessions and the Extended Self," *Journal of Consumer Research*, 1988, *15* (2), 139–168.

Berinsky, Adam J., *In Time of War: Understanding American Public Opinion from World War II to Iraq*, University of Chicago Press, October 2009.

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

Branscombe, Nyla R. and et al., "The Context and Content of Social Identity Threat," in "Social Identity: Context, Commitment, Content," Oxford, England: Blackwell Publishers Inc., 1999, pp. 35–58.

Bronnenberg, Bart J. and Jean-Pierre Dubé, "The Formation of Consumer Brand Preferences," *Annual Review of Economics*, 2017, *9* (1).

_ , Michael W. Kruger, and Carl F. Mela, "Database Paper: The IRI Marketing Data Set," *Marketing Science*, 2008, *27* (4), 745–748.

Carney, Dana R., John T. Jost, Samuel D. Gosling, and Jeff Potter, "The Secret Lives of Liberals and Conservatives: Personality Profiles, Interaction Styles, and the Things They Leave Behind," *Political Psychology*, 2008, *29* (6), 807–840.

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

Cutright, Keisha M., Eugenia C. Wu, Jillian C. Banfield, Aaron C. Kay, and Gavan J. Fitzsimons, "When Your World Must Be Defended: Choosing Products to Justify the System," *Journal of Consumer Research*, 2011, *38* (1), 62–77.

Davies, Paul G., Claude M. Steele, and Hazel Rose Markus, "A Nation Challenged: The Impact of Foreign Threat on America's Tolerance for Diversity," *Journal of Personality and Social Psychology*, 2008, *95* (2), 308–318.

Fearon, James D., "Domestic Political Audiences and the Escalation of International Disputes," *American Political Science Review*, 1994, *88* (3), 577–592.

Friese, Malte and Wilhelm Hofmann, "What Would You have as a Last Supper? Thoughts About Death Influence Evaluation and Consumption of Food Products," *Journal of Experimental Social Psychology*, September 2008, *44* (5), 1388–1394.

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

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

_ , "Evaluating Claims of Social Connection to International Conflict Casualties," *International Interactions*, August 2009, *35* (3), 352–364.

— **and Gary M. Segura**, "Race, Casualties, and Opinion in the Vietnam War," *Journal of Politics*, 2000, *62* (1), 115–146.

Gelpi, Christopher, Peter D. Feaver, and Jason Reifler, *Paying the Human Costs of War: American Public Opinion and Casualties in Military Conflicts*, Princeton University Press, February 2009.

Gerber, Alan S. and et al., "How Large and Long-lasting Are the Persuasive Effects of Televised Campaign Ads? Results from a Randomized Field Experiment," *American Political Science Review*, February 2011, *105* (1), 135–150.

Gershkoff, Amy and Shana Kushner, "Shaping Public Opinion: The 9/11-Iraq Connection in the Bush Administration's Rhetoric," *Perspectives on Politics*, 2005, *3* (3), 525–537.

Gourevitch, Peter, "The Second Image Reversed: The International Sources of Domestic Politics," *International Organization*, 1978, *32* (4), 881–912.

Greenberg, Jeff and et al., "Terror Management and Tolerance: Does Mortality Salience Always Intensify Negative Reactions to Others who Threaten One's Worldview?," *Journal of Personality and Social Psychology*, 1992, *63* (2), 212–220.

— **, Sheldon Solomon, and Tom Pyszczynski**, "Terror Management Theory of Self-Esteem and Cultural Worldviews: Empirical Assessments and Conceptual Refinements," in Mark P. Zanna, ed., *Advances in Experimental Social Psychology*, Vol. 29, Academic Press, January 1997, pp. 61–139.

Grossman, Gene M. and Elhanan Helpman, "Protection for Sale," *American Economic Review*, 1994, *84* (4), 833–850.

Guisinger, Alexandra, *American Opinion on Trade: Preferences without Politics*, New York: Oxford University Press, 2017.

Gürhan-Canli, Zeynep and Durairaj Maheswaran, "Cultural Variations in Country of Origin Effects," *Journal of Marketing Research*, August 2000, *37* (3), 309–317.

Hafner-Burton, Emilie M., Stephan Haggard, David A. Lake, and David G. Victor, "The Behavioral Revolution and International Relations," *International Organization*, April 2017, *71* (S1), S1–S31.

Hainmueller, Jens and Michael J. Hiscox, "Attitudes toward Highly Skilled and Low-Skilled Immigration: Evidence from a Survey Experiment," *American Political Science Review*, 2010, *104* (1), 61–84.

Hawkes, William J. and Frank W. Piotrowski, *Using Scanner Data to Improve the Quality of Measurement in the Consumer Price Index*, University of Chicago Press, January

Herrmann, Richard K., "How Attachments to the Nation Shape Beliefs About the World: A Theory of Motivated Reasoning," *International Organization*, April 2017, *71* (S1), S61–S84.

Hetherington, Marc and Elizabeth Suhay, "Authoritarianism, Threat, and Americans' Support for the War on Terror," *American Journal of Political Science*, 2011, *55* (3), 546–560.

Huddy, Leonie and Nadia Khatib, "American Patriotism, National Identity, and Political Involvement," *American Journal of Political Science*, 2007, *51* (1), 63–77.

_ , Stanley Feldman, Charles Taber, and Gallya Lahav, "Threat, Anxiety, and Support of Antiterrorism Policies," *American Journal of Political Science*, 2005, *49* (3), 593–608.

_ , _ , Theresa Capelos, and Colin Provost, "The Consequences of Terrorism: Disentangling the Effects of Personal and National Threat," *Political Psychology*, 2002, *23* (3), 485–509.

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

Jost, John T., "The marketplace of ideology: "Elective affinities" in political psychology and their implications for consumer behavior," *Journal of Consumer Psychology*, 2017, *27* (4), 502–520.

Kahn, Barbara E. and David C. Schmittlein, "Shopping Trip Behavior: An Empirical Investigation," *Marketing Letters*, 1989, *1* (1), 55–69.

Karol, David and Edward Miguel, "The Electoral Cost of War: Iraq Casualties and the 2004 U.S. Presidential Election," *The Journal of Politics*, August 2007, *69* (3), 633–648.

Kertzer, Joshua D. and Dustin Tingley, "Political Psychology in International Relations: Beyond the Paradigms," *Annual Review of Political Science*, 2018, *21* (1), 319–339.

_ and Thomas Zeitzoff, "A Bottom-Up Theory of Public Opinion about Foreign Policy," *American Journal of Political Science*, 2017, *61* (3), 543–558.

Khan, Romana, Kanishka Misra, and Vishal Singh, "Ideology and Brand Consumption," *Psychological Science*, 2013, *24* (3), 326–333.

Klein, Jill, "Us Versus Them, or Us Versus Everyone? Delineating Consumer Aversion to Foreign Goods," *Journal of International Business Studies*, 2002, *33*, 345–363.

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

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

Kriner, Douglas and Francis Shen, "Limited War and American Political Engagement," *The Journal of Politics*, October 2009, *71* (4), 1514–1529.

Kriner, Douglas L. and Francis X. Shen, *The Casualty Gap: The Causes and Consequences of American Wartime Inequalities*, Oxford, England: Oxford University Press, 2010.

Levay, Kevin E., Jeremy Freese, and James N. Druckman, "The Demographic and Political Composition of Mechanical Turk Samples," *SAGE Open*, 2016, *6* (1), 1–17.

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

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

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

_ **and Cathy Yi Chen**, "Nation Equity: Incidental Emotions in Country-of-Origin Effects," *Journal of Consumer Research*, 2006, *33* (3), 370–376.

Mansfield, Edward D. and Diana C. Mutz, "US versus Them: Mass Attitudes toward Offshore Outsourcing," *World Politics*, 2013, *65* (4), 571–608.

Margalit, Yotam, "Political Responses to Economic Shocks," *Annual Review of Political Science*, 2019, *22*, 277–295.

Martin, Brett A.S., Michael Shyue Wai Lee, and Charlotte Lacey, "Countering Negative Country of Origin Effects Using Imagery Processing," *Journal of Consumer Behaviour*, 2011, *10* (2), 80–92.

Mutz, Diana C., "Status Threat, Not Economic Hardship, Explains the 2016 Presidential Vote," *Proceedings of the National Academy of Sciences*, May 2018, *115* (19), E4330–E4339.

Myrick, Rachel, "Do External Threats Unite or Divide? Security Crises, Rivalries, and Polarization in American Foreign Policy," *International Organization*, 2021.

Naoi, Megumi, "Survey Experiments in International Political Economy: What We (Don't) Know About the Backlash Against Globalization," *Annual Review of Political Science*, 2020, *23* (1), 333–356.

Nelson, Lori J and et al., "General and Personal Mortality Salience and Nationalistic Bias," *Personality and Social Psychology Bulletin*, 1997, *23* (8), 884–892.

Rindfleisch, Aric, James E. Burroughs, and Nancy Wong, "The Safety of Objects: Materialism, Existential Insecurity, and Brand Connection," *Journal of Consumer Research*, June 2009, *36* (1), 1–16.

Samiee, Saeed, Terence A. Shimp, and Subhash Sharma, "Brand Origin Recognition Accuracy: Its Antecedents and Consumers' Cognitive Limitations," *Journal of International Business Studies*, July 2005, *36* (4).

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

Shachar, Ron, Tülin Erdem, Keisha M. Cutright, and Gavan J. Fitzsimons, "Brands: The Opiate of the Nonreligious Masses?," *Marketing Science*, 2011, *30* (1), 92–110.

Sharma, Subhash, Terence A. Shimp, and Jeongshin Shin, "Consumer Ethnocentrism: A Test of Antecedents and Moderators," *Journal of the Academy of Marketing Science*, December 1995, *23* (1), 26–37.

Shepherd, Steven, Tanya L. Chartrand, and Gavan J. Fitzsimons, "When Brands Reflect Our Ideal World: The Values and Brand Preferences of Consumers Who Support versus Reject Society's Dominant Ideology," *Journal of Consumer Research*, 2015, *42* (1), 76–92.

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

Sorensen, Herb, Svetlana Bogomolova, Katherine Anderson, Giang Trinh, Anne Sharp, Rachel Kennedy, Bill Page, and Malcolm Wright, "Fundamental Patterns of In-Store Shopper Behavior," *Journal of Retailing and Consumer Services*, July 2017, *37*, 182–194.

Steenkamp, Jan-Benedict E.M., Rajeev Batra, and Dana L. Alden, "How Perceived Brand Globalness Creates Brand Value," *Journal of International Business Studies*, 2003, *34* (1), 53–65.

Tajfel, Henri, *Human Groups and Social Categories: Studies in Social Psychology*, Cambridge: Cambridge University Press, 1981.

_ **and John C. Turner**, "An Integrative Theory of Intergroup Conflict," in William G. Austin and Stephen Worchel, eds., *The Social Psychology of Intergroup Relations*, Monterey, CA: Brooks/Cole Publishing Company, 1979, pp. 33–47.

Terry, Deborah J., Michael A. Hogg, and Katherine M. White, "The Theory of Planned Behavior: Self-Identity, Social Identity, and Group Norms," *British Journal of Social Psychology*, 1999, *38*, 225–244.

Valentino, Nicholas A., Vincent L. Hutchings, and Dmitri Williams, "The Impact of Political Advertising on Knowledge, Internet Information Seeking and Candidate Preference," *Journal of Communication*, January 2004, *54* (2), 337–354.

Verlegh, Peeter W.J. and Jan-Benedic E.M. Steenkamp, "A Review and Meta-Analysis of Country-of-Origin Research," *Journal of Economic Psychology*, 1999, *20*, 521–546.

White, Katherine and Darren W. Dahl, "To Be or Not Be? The Influence of Dissociative Reference Groups on Consumer Preferences," *Journal of Consumer Psychology*, January 2006, *16* (4), 404–414.

_ **and** _ , "Are All Out-Groups Created Equal? Consumer Identity and Dissociative Influence," *Journal of Consumer Research*, December 2007, *34* (4), 525–536.

Yang, Sha, Chen Yuxin, and Greg M. Allenby, "Bayesian Analysis of Simultaneous Demand and Supply," *Quantitative Marketing and Economics*, 2003, *1*, 251–275.

A Appendix: War on Aisle 5

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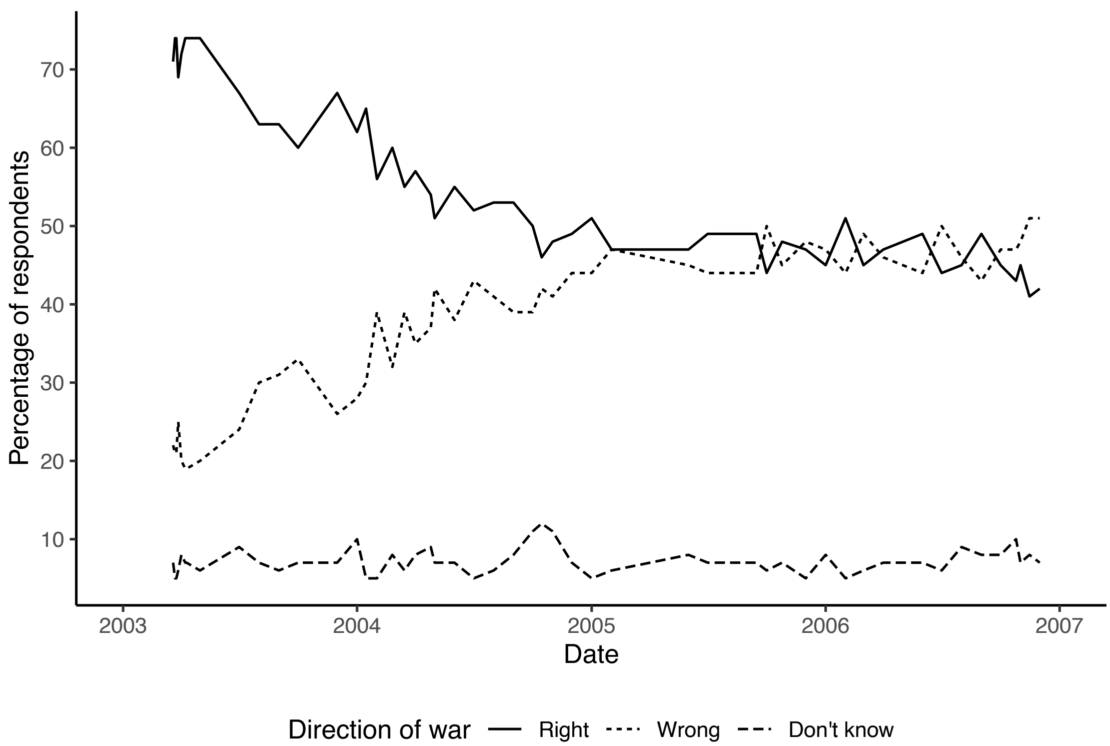
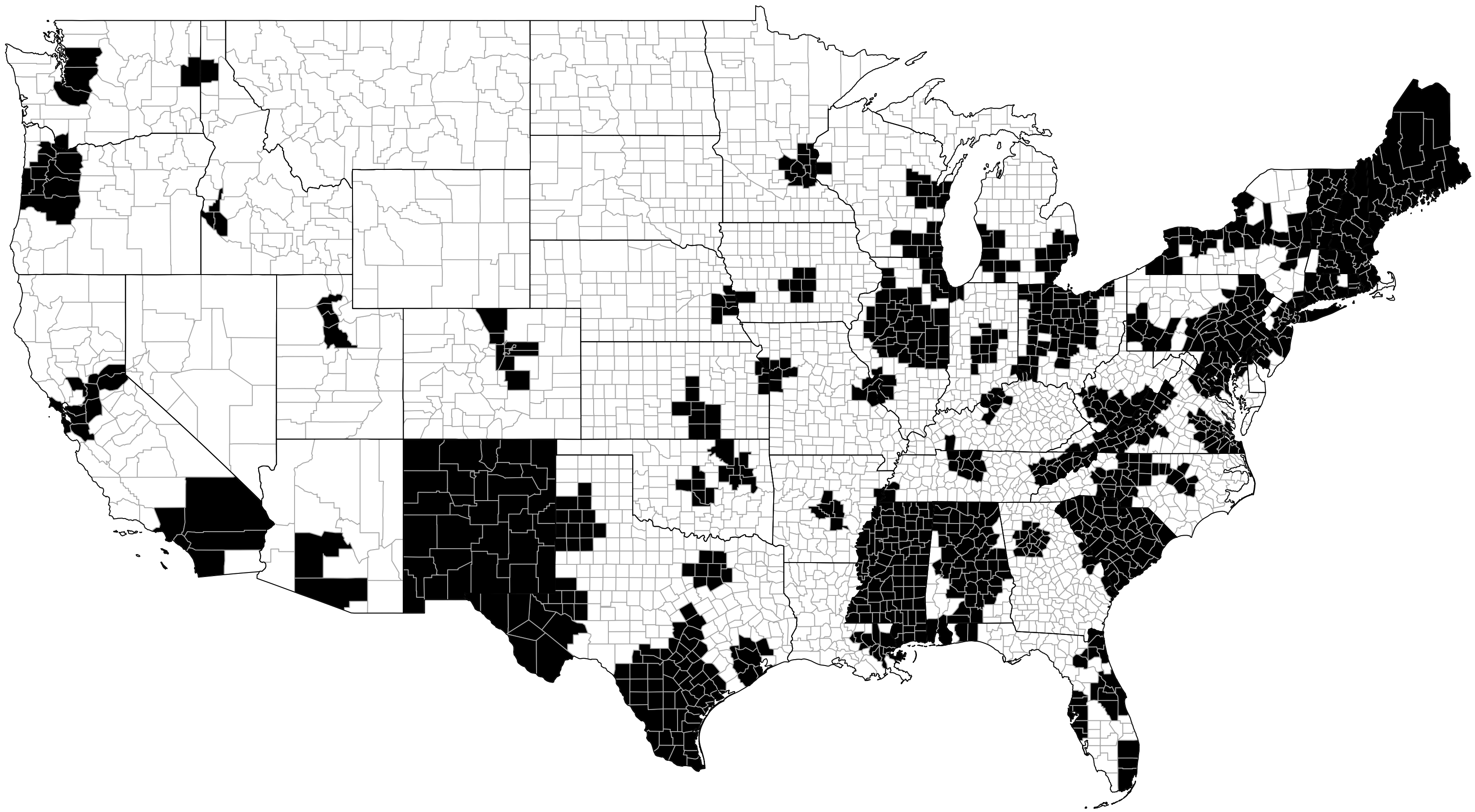
Appendix Figure A.1: American Support for Iraq War Declined, 2003-2006



Data Source: Pew Research Center.

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 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

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 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: Weekly Casualties and American Brand Share, 2003-2006

	$\Delta ShareYear\text{-}2001_{ijkt}$ (%)			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$LocalCasualty_{jt} * AmericanScore_i$	0.014**	0.012**	0.011**	0.030***
	(0.006)	(0.005)	(0.006)	(0.006)
$NationalCasualties_t * AmericanScore_i$	0.0001	−0.0002**	0.001***	−0.0001
	(0.0001)	(0.0001)	(0.0001)	(0.0002)
$LocalCasualty_{jt}$	0.025	−0.095***	−0.050*	−0.152***
	(0.029)	(0.023)	(0.026)	(0.026)
$NationalCasualties_t$	0.001	−0.001	−0.001**	0.005***
	(0.0004)	(0.0004)	(0.001)	(0.001)
$AmericanScore_i$	−0.013***	−0.052***	−0.122***	−0.091***
	(0.003)	(0.003)	(0.003)	(0.004)
$HomePrice_{jt}$	0.004***	0.002***	−0.002***	−0.003***
	(0.0003)	(0.0003)	(0.0003)	(0.0002)
$HomePrice_{jt} * AmericanScore_i$	−0.001***	−0.0005***	0.0002***	0.0005***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
$Enlistment_j$	−0.105***	−0.261***	−0.331***	−0.348***
	(0.027)	(0.029)	(0.032)	(0.033)
$Enlistment_j * AmericanScore_i$	−0.013**	0.027***	0.059***	0.060***
	(0.006)	(0.006)	(0.007)	(0.007)
$Population_{j2000}$	0.00002***	0.00003***	0.00002***	0.00002***
	(0.00000)	(0.00000)	(0.00000)	(0.00000)
$\Delta PriceYear\text{-}2001_{ijkt}$	−0.053***	−0.052***	−0.040***	−0.019***
	(0.001)	(0.001)	(0.001)	(0.001)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.900***	0.857***	0.872***	0.968***
	(0.001)	(0.001)	(0.001)	(0.001)
$Intercept$	−0.042***	0.127***	0.433***	0.358***
	(0.012)	(0.014)	(0.016)	(0.019)
Observations	6,715,772	6,344,222	5,756,986	5,533,301

Note: *p<0.1; **p<0.05; ***p<0.01. All models estimated with OLS.

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

	Δ*ShareYear-2001*$_{ijkt}$ (%)			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$LocalCasualty_{jt-1} * AmericanScore_i$	0.02178***	0.00905*	0.01544***	0.02996***
	(0.00636)	(0.00504)	(0.00577)	(0.00584)
$NationalCasualties_{t-1} * AmericanScore_i$	0.00011	−0.00068***	0.00092***	0.00073***
	(0.00009)	(0.00009)	(0.00012)	(0.00019)
$LocalCasualty_{jt-1}$	0.03606	−0.10498***	−0.07919***	−0.14764***
	(0.02922)	(0.02312)	(0.02647)	(0.02677)
$NationalCasualties_{t-1}$	−0.00103**	0.00173***	−0.00417***	0.00112
	(0.00042)	(0.00042)	(0.00054)	(0.00086)
$AmericanScore_i$	−0.01207***	−0.04358***	−0.12637***	−0.10291***
	(0.00267)	(0.00304)	(0.00344)	(0.00419)
$HomePrice_{jt}$	0.00402***	0.00164***	−0.00166***	−0.00278***
	(0.00030)	(0.00027)	(0.00025)	(0.00025)
$HomePrice_{jt} * AmericanScore_i$	−0.00113***	−0.00047***	0.00024***	0.00045***
	(0.00007)	(0.00006)	(0.00005)	(0.00005)
$Enlistment_j$	−0.09905***	−0.25925***	−0.33229***	−0.34049***
	(0.02774)	(0.02920)	(0.03226)	(0.03291)
$Enlistment_j * AmericanScore_i$	−0.01428**	0.02660***	0.05905***	0.05887***
	(0.00598)	(0.00634)	(0.00700)	(0.00717)
$Population_{j2000}$	0.00002***	0.00003***	0.00002***	0.00002***
	(0.000001)	(0.000001)	(0.000001)	(0.000002)
Δ*PriceYear-2001*$_{ijkt}$	−0.05347***	−0.05229***	−0.03981***	−0.01923***
	(0.00101)	(0.00099)	(0.00099)	(0.00076)
Δ*VariantsYear-2001*$_{ijkt}$	0.89919***	0.85594***	0.87459***	0.96859***
	(0.00102)	(0.00091)	(0.00094)	(0.00104)
Δ*VariantsYear-2001*$_{ijkt}$	−0.03128**	0.08639***	0.48636***	0.42120***
	(0.01231)	(0.01393)	(0.01577)	(0.01915)
Observations	6,569,912	6,206,893	5,630,408	5,413,818

Note: *p<0.1; **p<0.05; ***p<0.01. All models estimated with OLS.

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)
$ln(CumulCasualties_{jt}) * AmericanScore_i$	0.021***	0.018***	0.012***	0.010***
	(0.002)	(0.001)	(0.001)	(0.001)
$ln(CumulNatCasualties_{jt}) * AmericanScore_i$	−0.004***	−0.088***	0.052***	0.090***
	(0.0005)	(0.004)	(0.009)	(0.015)
$ln(CumulCasualties_{jt})$	0.006	−0.097***	−0.081***	−0.060***
	(0.010)	(0.007)	(0.007)	(0.007)
$ln(CumulNatCasualties_{jt})$	0.004*	0.338***	0.286***	0.724***
	(0.002)	(0.019)	(0.043)	(0.066)
$AmericanScore_i$	0.007**	0.544***	−0.506***	−0.809***
	(0.003)	(0.028)	(0.071)	(0.115)
$HomePrice_{jt}$	0.004***	0.003***	−0.001**	−0.002***
	(0.0003)	(0.0003)	(0.0003)	(0.0003)
$HomePrice_{jt} * AmericanScore_i$	−0.001***	−0.001***	0.00002	0.0003***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
$Enlistment_j$	−0.062**	−0.139***	−0.211***	−0.255***
	(0.028)	(0.030)	(0.034)	(0.035)
$Enlistment_j * AmericanScore_i$	−0.026***	0.0004	0.036***	0.042***
	(0.006)	(0.007)	(0.007)	(0.008)
$Population_{j2000}$	0.00001***	0.00004***	0.00003***	0.00002***
	(0.00000)	(0.00000)	(0.00000)	(0.00000)
$\Delta PriceYear\text{-}2001_{ijkt}$ (2001-Year)	−0.053***	−0.052***	−0.040***	−0.019***
	(0.001)	(0.001)	(0.001)	(0.001)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.901***	0.857***	0.872***	0.968***
	(0.001)	(0.001)	(0.001)	(0.001)
$Intercept$	−0.062***	−2.187***	−1.712***	−5.248***
	(0.015)	(0.128)	(0.323)	(0.524)
Observations	6,715,772	6,344,222	5,756,986	5,533,301

Note: *p<0.1; **p<0.05; ***p<0.01. All models estimated with OLS.

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)
$CumulCasualties_{jt}/Pop_{j2000} * AmericanScore_i$	1.682***	0.577**	0.838***	0.483***
	(0.607)	(0.277)	(0.225)	(0.178)
$ln(CumulNatCasualties_{jt}) * AmericanScore_i$	−0.003***	−0.079***	0.055***	0.095***
	(0.0005)	(0.004)	(0.009)	(0.015)
$CumulCasualties_{jt}/Population_{j2000}$	0.723	−2.747**	−9.463***	−7.052***
	(2.824)	(1.273)	(1.033)	(0.816)
$ln(CumulNatCasualties_{jt})$	0.004**	0.286***	0.275***	0.726***
	(0.002)	(0.019)	(0.043)	(0.066)
$AmericanScore_i$	−0.001	0.486***	−0.532***	−0.846***
	(0.003)	(0.028)	(0.071)	(0.115)
$HomePrice_{jt}$	0.005***	0.002***	−0.001***	−0.002***
	(0.0003)	(0.0003)	(0.0002)	(0.0002)
$HomePrice_{jt} * AmericanScore_i$	−0.001***	−0.0004***	0.0002***	0.001***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
$Enlistment_j$	−0.020	−0.176***	−0.291***	−0.309***
	(0.027)	(0.028)	(0.031)	(0.032)
$Enlistment_j * AmericanScore_i$	−0.011*	0.029***	0.060***	0.064***
	(0.006)	(0.006)	(0.007)	(0.007)
$\Delta PriceYear\text{-}2001_{ijkt}$	−0.054***	−0.053***	−0.040***	−0.019***
	(0.001)	(0.001)	(0.001)	(0.001)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.901***	0.857***	0.872***	0.968***
	(0.001)	(0.001)	(0.001)	(0.001)
$Intercept$	−0.071***	−1.847***	−1.615***	−5.256***
	(0.015)	(0.127)	(0.323)	(0.524)
Observations	6,715,772	6,344,222	5,756,986	5,533,301

Note: *p<0.1; **p<0.05; ***p<0.01. All models estimated using OLS.

Appendix Table A.9: Weekly Casualties and American Brand Share, 2003-2006 - Dichotomized AmericanScore$_i$

	Δ*ShareYear-2001$_{ijkt}$* (%)			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$LocalCasualty_{jt} * American_i$	0.14014***	0.11466***	0.15017***	0.25819***
	(0.03758)	(0.02976)	(0.03329)	(0.03307)
$NationalCasualties_t * American_i$	0.00026	−0.00020	0.00088	0.00036
	(0.00054)	(0.00055)	(0.00069)	(0.00108)
$LocalCasualty_{jt}$	−0.02087	−0.10074***	−0.07476***	−0.16549***
	(0.03064)	(0.02421)	(0.02716)	(0.02694)
$NationalCasualties_t$	0.00255***	−0.00083*	0.00162***	0.00316***
	(0.00044)	(0.00045)	(0.00056)	(0.00087)
$American_i$	0.07769***	−0.12961***	−0.21583***	−0.15159***
	(0.01515)	(0.01729)	(0.01924)	(0.02338)
$HomePrice_{jt}$	0.00456***	0.00233***	0.00140***	−0.00037
	(0.00034)	(0.00030)	(0.00027)	(0.00027)
$HomePrice_{jt} * American_i$	−0.00152***	−0.00089***	−0.00053***	−0.00020***
	(0.00006)	(0.00006)	(0.00005)	(0.00005)
$Enlistment_j$	0.20404***	0.23436***	0.63248***	0.45213***
	(0.03073)	(0.03193)	(0.03467)	(0.03471)
$Enlistment_j * American_i$	−0.07238***	−0.06805***	−0.11080***	−0.09051***
	(0.00562)	(0.00590)	(0.00640)	(0.00644)
$Population_{j2000}$	0.00003***	0.00005***	0.00003***	0.00002***
	(0.000002)	(0.000002)	(0.000002)	(0.000002)
$\Delta PriceYear\text{-}2001_{ijkt}$	−0.02827***	−0.02791***	−0.01064***	−0.00296***
	(0.00117)	(0.00118)	(0.00118)	(0.00109)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.92611***	0.81387***	0.79256***	0.97539***
	(0.00151)	(0.00121)	(0.00129)	(0.00156)
$Intercept$	−0.24522***	−0.16372***	−0.25395***	−0.19642***
	(0.01262)	(0.01425)	(0.01586)	(0.01910)
Observations	3,243,529	3,056,611	2,778,396	2,675,991

Note: *p<0.1; **p<0.05; ***p<0.01. Sample limited to $AmericanScore_i < 2$ and $AmericanScore_i > 5$. $AmericanScore_i < 2$ recoded to 0 and $AmericanScore_i > 5$ recoded to 1 to create $American_i$.

Appendix Table A.10: Cumulative Casualties and American Brand Share, 2003-2006 - Dichotomized AmericanScore$_\mathbf{i}$

	Change in market share for American brands			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * American_i$	0.16473***	0.16141***	0.19589***	0.18499***
	(0.01201)	(0.00837)	(0.00799)	(0.00793)
$ln(CumulNatCasualties_{jt}) * American_i$	−0.04158***	−0.52419***	0.23901***	0.39819***
	(0.00288)	(0.02426)	(0.05484)	(0.08357)
$ln(CumulCasualties_{jt})$	−0.02850***	−0.10609***	−0.13187***	−0.12671***
	(0.01057)	(0.00783)	(0.00755)	(0.00758)
$ln(CumulNatCasualties_{jt})$	0.01914***	0.22594***	0.41193***	0.73165***
	(0.00238)	(0.01983)	(0.04453)	(0.06754)
$American_i$	0.26479***	3.42844***	−2.09521***	−3.43855***
	(0.01890)	(0.16634)	(0.41390)	(0.66158)
$HomePrice_{jt}$	0.00514***	0.00387***	0.00341***	0.00149***
	(0.00034)	(0.00031)	(0.00029)	(0.00028)
$HomePrice_{jt} * American_i$	−0.00168***	−0.00123***	−0.00103***	−0.00063***
	(0.00006)	(0.00006)	(0.00005)	(0.00005)
$Enlistment_j$	0.25951***	0.38530***	0.86713***	0.67591***
	(0.03114)	(0.03311)	(0.03640)	(0.03651)
$Enlistment_j * American_i$	−0.08709***	−0.10247***	−0.16435***	−0.14001***
	(0.00574)	(0.00617)	(0.00677)	(0.00682)
$Population_{j2000}$	0.00002***	0.00005***	0.00004***	0.00003***
	(0.000002)	(0.000002)	(0.000003)	(0.000003)
$\Delta PriceYear\text{-}2001_{ijkt}$	−0.02838***	−0.02819***	−0.01062***	−0.00295***
	(0.00117)	(0.00118)	(0.00118)	(0.00109)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.92677***	0.81481***	0.79315***	0.97539***
	(0.00151)	(0.00121)	(0.00129)	(0.00156)
$Intercept$	−0.30532***	−1.70722***	−3.27046***	−5.84105***
	(0.01565)	(0.13581)	(0.33593)	(0.53449)
Observations	3,243,529	3,056,611	2,778,396	2,675,991

Note: *p<0.1; **p<0.05; ***p<0.01. Sample limited to $AmericanScore_i < 2$ and $AmericanScore_i > 5$. $AmericanScore_i < 2$ recoded to 0 and $AmericanScore_i > 5$ recoded to 1 to create $American_i$.

Appendix Table A.11: Baseline Demographic Propensity for American Brands - 2001

	$Share2001_{ijkt}$ (%)
	(1)
$AmericanScore_i * Bush_{j2000}$	0.00036^{***}
	(0.00013)
$AmericanScore_i * SkilledOcc_{j2000}$	0.00405^{***}
	(0.00019)
$AmericanScore_i * ArmedForces_{j2000}$	0.00603^{***}
	(0.00065)
$AmericanScore_i * NativeBorn_{j2000}$	0.00008
	(0.00017)
$AmericanScore_i * Black_{j2000}$	-0.00028^{**}
	(0.00011)
$AmericanScore_i$	-0.08147^{***}
	(0.01535)
Observations	3,467,761
Controls	✓
Category FEs	✓
Store FEs	✓
Week FEs	✓

Note: *p<0.1; **p<0.05; ***p<0.01. Model estimated with OLS. Controls include $Price2001_{ijkt}$ and $Variants2001_{ijkt}$. Sample limited to weeks in 2001 prior to September 11. Stores in counties with populations two standard deviations above or below the mean excluded.

Appendix Table A.12: Cumulative Casualties and Coalition of the Willing Brand Share, 2003-2006

	$\Delta ShareYear\text{-}2001_{ijkt}$			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * CoalitionScore_i$	-0.02541^{***}	-0.01616^{***}	-0.00483^{**}	-0.00844^{***}
	(0.00312)	(0.00231)	(0.00225)	(0.00237)
$ln(CumulNatCasualties_{jt}) * CoalitionScore_i$	0.00287^{***}	0.03994^{***}	-0.04509^{***}	-0.04148^{*}
	(0.00075)	(0.00658)	(0.01514)	(0.02458)
$ln(CumulCasualties_{jt})$	0.08441^{***}	-0.11066^{***}	-0.14052^{***}	-0.02233^{***}
	(0.01012)	(0.00788)	(0.00771)	(0.00822)
$ln(CumulNatCasualties_{jt})$	-0.01208^{***}	0.30525^{***}	1.41098^{***}	1.13861^{***}
	(0.00222)	(0.01943)	(0.04452)	(0.07221)
$CoalitionScore_i$	0.00746	-0.22283^{***}	0.41042^{***}	0.41848^{**}
	(0.00513)	(0.04516)	(0.11432)	(0.19461)
$HomePrice_{jt}$	-0.00330^{***}	-0.00268^{***}	-0.00031	-0.00094^{***}
	(0.00032)	(0.00031)	(0.00029)	(0.00029)
$HomePrice_{jt} * CoalitionScore_i$	-0.00020^{*}	-0.00034^{***}	-0.00068^{***}	-0.00064^{***}
	(0.00010)	(0.00010)	(0.00010)	(0.00010)
$Enlistment_j$	-0.34477^{***}	-0.07142^{**}	-0.00530	0.11432^{***}
	(0.02837)	(0.03180)	(0.03551)	(0.03803)
$Enlistment_j * CoalitionScore_i$	0.00846	-0.02342^{**}	-0.03673^{***}	-0.06082^{***}
	(0.00944)	(0.01073)	(0.01200)	(0.01285)
$Population_{j2000}$	0.00008^{***}	0.00010^{***}	0.00010^{***}	0.00011^{***}
	(0.000002)	(0.000002)	(0.000003)	(0.000003)
$\Delta PriceYear\text{-}2001_{ijkt}$	-0.08603^{***}	-0.06060^{***}	-0.05599^{***}	-0.02745^{***}
	(0.00158)	(0.00149)	(0.00147)	(0.00112)
$\Delta VariantsYear\text{-}2001_{ijkt}$	0.76665^{***}	0.74392^{***}	0.72880^{***}	0.68107^{***}
	(0.00092)	(0.00092)	(0.00090)	(0.00089)
$Intercept$	-0.10034^{***}	-2.28969^{***}	-10.55783^{***}	-9.10191^{***}
	(0.01506)	(0.13320)	(0.33587)	(0.57155)
Observations	6,573,689	5,968,261	5,472,115	5,056,532

Note: *p<0.1; **p<0.05; ***p<0.01. All models estimated using OLS.

Appendix Table A.13: Cumulative Casualties and France/Germany Brand Share, 2003-2006

	Δ*ShareYear-2001*$_{ijkt}$			
	2003	2004	2005	2006
	(1)	(2)	(3)	(4)
$ln(CumulCasualties_{jt}) * France + GermScore_i$	-0.02369^{***}	-0.00806^{***}	-0.00718^{***}	-0.00463^{*}
	(0.00338)	(0.00249)	(0.00243)	(0.00252)
$ln(CumulNatCasualties_{jt}) * France + GermScore_i$	0.00467^{***}	0.03935^{***}	0.11508^{***}	0.04034
	(0.00081)	(0.00702)	(0.01626)	(0.02598)
$ln(CumulCasualties_{jt})$	0.05646^{***}	-0.11765^{***}	-0.11885^{***}	-0.03885^{***}
	(0.01119)	(0.00875)	(0.00860)	(0.00906)
$ln(CumulNatCasualties_{jt})$	-0.02137^{***}	0.33990^{***}	0.99000^{***}	1.14468^{***}
	(0.00241)	(0.02090)	(0.04795)	(0.07651)
$France + GermScore_i$	-0.09513^{***}	-0.33963^{***}	-0.93750^{***}	-0.38469^{*}
	(0.00551)	(0.04821)	(0.12275)	(0.20568)
$HomePrice_{jt}$	-0.00246^{***}	-0.00244^{***}	0.00002	-0.00036
	(0.00034)	(0.00033)	(0.00031)	(0.00031)
$HomePrice_{jt} * France + GermanyScore_i$	0.00045^{***}	0.00042^{***}	0.00026^{**}	0.00018^{*}
	(0.00011)	(0.00011)	(0.00010)	(0.00011)
$Enlistment_j$	-0.24638^{***}	-0.13918^{***}	-0.04734	-0.00245
	(0.03086)	(0.03424)	(0.03831)	(0.04039)
$Enlistment_j * France + GermanyScore_i$	-0.04167^{***}	-0.03676^{***}	-0.02888^{**}	-0.02752^{**}
	(0.01029)	(0.01161)	(0.01300)	(0.01369)
$Population_{j2000}$	0.00007^{***}	0.00009^{***}	0.00009^{***}	0.00009^{***}
	(0.000002)	(0.000003)	(0.000003)	(0.000003)
Δ*PriceYear-2001*$_{ijkt}$	-0.11624^{***}	-0.09472^{***}	-0.09314^{***}	-0.06908^{***}
	(0.00190)	(0.00185)	(0.00186)	(0.00164)
Δ*VariantsYear-2001*$_{ijkt}$	0.78076^{***}	0.76209^{***}	0.75684^{***}	0.73213^{***}
	(0.00116)	(0.00114)	(0.00115)	(0.00116)
Intercept	0.07102^{***}	-2.27963^{***}	-7.15117^{***}	-8.82601^{***}
	(0.01628)	(0.14320)	(0.36162)	(0.60543)
Observations	4,770,144	4,339,951	3,966,649	3,662,239

Note: *p<0.1; **p<0.05; ***p<0.01. All models estimated using OLS.

A.1 External Validation of AmericanScore$_b$

As an external validation exercise, we verify that American brands indeed symbolize American national identity. The brands that score high (low) on $AmericanScore_b$ should be strong (weak) symbols of America. We measure brands' America symbolism using a three-item scale from Steenkamp et al (2003): "To me, this brand is a symbol of America; I associate this brand with things that are American; To me, this brand represents American values." The rating options ranged from 1 to 7, with 1 meaning "I completely disagree with this statement," and 7 meaning "I completely agree with this statement." As such, we expect a correlation between our $AmericanScore_b$ and the American Symbolism scales.

Given the massive number of brands (8,644) already evaluated with the $AmericanScore_b$ scale, we decided to sample a selection of them (40 brands) to test the correlation between the $AmericanScore_b$ and the American Symbolism scales. We randomly selected five different product categories (beer, laundry detergent, toothpaste, spaghetti sauce, and frozen dinner). For each one of these categories, we needed to include brands belonging to different levels of $AmericanScore_b$. As such, we randomly selected two brands with an American Score of zero, three, five, and seven. Summarizing, the selection of these 40 brands was the result of two randomly selected brands for each of four levels of $AmericanScore_b$ in five different brand categories.

We assessed that evaluating 40 brands in a single session could generate survey fatigue. As such, we assigned participants to evaluate only 20 brands (instead of 40) in a random order. Each respondent evaluated a brand that scored zero, three, five, and seven on each of the five different brand categories. This generated a mix-model design, in which we had 20 brands, within-subjects (four $AmericanScore_b$ levels five brand categories), and two randomly generated brand lists, between-subjects. To exemplify, Respondent 1 evaluated 20 brands (one brand for four different $AmericanScore_b$ levels five brand categories—what we called "List A"). Respondent 2 evaluated the other 20 brands (one brand for four different $AmericanScore_b$ levels five brand categories—what we called "List B"). Respondent 3 evaluated List A, Respondent 4 evaluated List B, and so on. Table 2 shows the 40 brands used and identifies to which of the two lists it belonged. 400 US-based participants from an online pool participated in this study in exchange for money.

Appendix Table A.14: Mean Values of American Symbolism, Test Brands

Category	List	Brand Name	American Score	American Symbolism Scale (mean)	American Symbolism Scale (95% CI)
beer	List A	Guinness	0	2.82	2.57, 3.06
beer	List B	Tequiza	0	2.25	2.07, 2.43
beer	List A	Sierra Blanca	3	2.70	2.50, 2.91
beer	List B	Molson	3	2.99	2.78, 3.20
beer	List A	Red Wolf	5	3.81	3.59, 4.03
beer	List B	Brooklyn Brewery	5	5.09	4.87, 5.31
beer	List A	Budweiser	7	5.81	5.63, 5.99
beer	List B	Great Lakes Brewing	7	5.18	4.98, 5.38
laundry detergent	List A	Blanca Nieves	0	2.41	2.21, 2.60
laundry detergent	List B	Paloma	0	2.81	2.62, 2.99
laundry detergent	List A	Ariel	3	3.18	2.97, 3.39
laundry detergent	List B	Citra Suds	3	3.11	2.91, 3.31
laundry detergent	List A	Ajax	5	4.69	4.46, 4.91
laundry detergent	List B	Method	5	4.08	3.86, 4.30
laundry detergent	List A	Arm & Hammer	7	5.56	5.39, 5.73
laundry detergent	List B	Tide	7	5.52	5.33, 5.70
toothpaste	List A	Elgydium	0	2.49	2.29, 2.69
toothpaste	List B	Dabur	0	2.17	2.00, 2.34
toothpaste	List A	Butler	3	3.40	3.18, 3.62
toothpaste	List B	Shane	3	3.28	3.06, 3.50
toothpaste	List A	Mentadent	5	3.65	3.43, 3.87

toothpaste	List B	Choice	5	3.73	3.51, 3.95
toothpaste	List A	Colgate	7	5.38	5.20, 5.56
toothpaste	List B	Crest	7	5.44	5.26, 5.62
spaghetti sauce	List A	Cucina Antica	0	2.41	2.21, 2.62
spaghetti sauce	List B	Anna Mario's	0	2.94	2.75, 3.14
spaghetti sauce	List A	Prego	3	3.41	3.19, 3.63
spaghetti sauce	List B	Roland	3	3.32	3.12, 3.52
spaghetti sauce	List A	Sonoma Gourmet	5	4.34	4.12, 4.57
spaghetti sauce	List B	Ragu	5	4.51	4.29, 4.74
spaghetti sauce	List A	California Seasonings	7	4.77	4.55, 4.99
spaghetti sauce	List B	Uncle Dave's	7	4.43	4.21, 4.65
frozen dinner	List A	Ajinomoto	0	2.27	2.07, 2.46
frozen dinner	List B	Gallina Blanca	0	2.31	2.13, 2.49
frozen dinner	List A	Michelina's Signature	3	3.65	3.43, 3.87
frozen dinner	List B	Bobby Salazars	3	3.01	2.82, 3.19
frozen dinner	List A	Healthy Choice	5	4.85	4.64, 5.06
frozen dinner	List B	Seeds of Change	5	3.62	3.41, 3.84
frozen dinner	List A	Boston Market	7	5.45	5.26, 5.64
frozen dinner	List B	Uncle Ben's	7	5.20	5.00, 5.41

Appendix Table A.15: Mean American Symbolism By $\mathbf{AmericanScore_b}$ Level

American Score	American Symbolism Scale (mean)	American Symbolism Scale (95% CI for mean)
0	2.49	2.43, 2.55
3	3.20	3.14, 3.27
5	4.24	4.17, 4.31
7	5.27	5.21, 5.34

We analyzed the Cronbach's alpha of the three items we used to measure American symbolism. Given that the alpha was high (.96), we decided to average these three items in an index ("American Symbolism scale"). We then performed a mix-model ANOVA with the two Lists (A and B) as the between-subjects independent variable, and the 20 brands evaluated by each person as the within-subjects dependent variable. The means of each brand are reported in Appendix Table A.14.

This mix-model ANOVA revealed a non-significant effect of the list (F(1, 398) = 1.30, p = .254, ηp^2 = .003) and significant main effect of the brands (different brand names were evaluated differently on the American Symbolism scale; F(19, 7,562) = 348.15, p < .001, ηp^2 = .467). The interaction between the two factors was significant (F(19, 7,562) = 19.63, p < .001, ηp^2 = .047). A significant interaction was unexpected but simply indicated that, within the same brand level, sometime the means were higher in List A and some other times were higher in List B.

Given that there was no general main effect of the of the Lists, we pooled the data and grouped the 20 brands of List A and 20 brands of List B together. We then ran a correlation between $AmericanScore_b$ (0, 3, 5, 7) and American Symbolism scale. The two scales were strongly and positively correlated, r(7,998) = .56, p < .001 (note that 7,998 degrees of freedom, n = 8,000, equals to 20 brands $*$ 400 people). Appendix Table A.15 shows how each subsequent level of $AmericanScore_b$ corresponded to a higher rating on the American Symbolism scale.

References

Gerber, Alan S. and et al. (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.

Steenkamp, Jan-Benedict E. M., Batra, Rajeev, Alden, Dana L. (2003) How Perceived Brand Globalness Creates Brand Value. *Journal of International Business Studies, 34*: 53-65.

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.