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The baseline-inflated multinomial logit model for international relations research

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Abstract

International relations scholars are often interested in nominal dependent variables, and commonly analyze such variables with multinomial logit (MNL) models that treat status quo outcomes (e.g. "peace") as a homogeneous baseline-choice category. However, recent studies of zero-inflation processes within international relations suggest that these baseline cases may often arise from two distinct sources. Specifically, some status quo responses are likely to correspond to observations that actively opted for this choice over all others, while the remaining status quo outcomes are likely to arise from observations that were unable to realistically register a non-status quo choice under any reasonable circumstances. Including both sets of responses within an MNL model's baseline category can bias the estimated effects of covariates, leading to faulty inferences. As a solution to this problem, this study considers a recently proposed baseline-inflated MNL (BIMNL) model that explicitly estimates and tests for heterogeneous populations of status quo observations. After discussing the model and its theoretical underpinnings, I demonstrate the BIMNL's utility through replications of two existing studies of political violence and cooperation within the areas of international relations and civil war.

Keywords

Civil war, multinomial logit, polytomous choice, territorial disputes, zero-inflation

The problem of zero-inflation in quantitative analyses of international relations is now widely recognized. In dyad-year studies of international conflict, for instance, the dyadic pairing of every country-year in the world produces a preponderance of peace dyads. Many of these peace dyads—such as, for example, a dyad involving Costa Rica and Bhutan—can be considered irrelevant to the extent that geographic distance and a lack of general political–economic interaction effectively prevent these two countries from ever experiencing the types of disagreements that typically lead to militarized conflict (Maoz and Russett,

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1993; Xiang, 2010). Given the numerous "irrelevant dyads" that accordingly arise in observational studies of international disputes, scholars have come to recognize that the inclusion of such cases can bias one's estimates of political interaction, often in indeterminate directions. On the other hand, dropping all potential irrelevant dyad observations based on some crude criteria—such as the truncation of all non-contiguous, non-major power dyads—typically entails the loss of upwards of 20% of all conflict events (Bennett and Stam, 2004; Braumoeller and Carson, 2011), comparably biasing one's estimates in an opposing manner (Lemke and Reed, 2001; Xiang, 2010).

In response to these competing challenges, international relations scholars have come to argue that keeping all potentially irrelevant cases in one's sample—while adjusting for each observation's inflation (i.e. irrelevance) probability empirically—is an ideal strategy. Drawing upon the zero-inflated count model literature to do so, seminal efforts in this area used zero-inflated Poisson or zero-inflated negative binomial models to simultaneously estimate a dyad's inflation-propensity alongside that same dyad's expected conflict frequency, such that the latter probabilities were estimated conditional on the former—often with separate but overlapping covariates affecting each process (e.g. Pevehouse, 2004). Building upon these findings, scholars have since also used zero-inflated count models, and their theoretical underpinnings, to empirically account for cross-national and subnational inflation-propensities within the *intrastate* conflict setting (e.g. Hegre et al., 2009; Holmes et al., 2007), as well as among terrorist events (Piazza, 2011). Moreover, while earlier empirical studies predominately sought to address zero-inflation within the count variable context (Benini and Moulton, 2004; Pevehouse, 2004), conflict scholars have more recently begun to apply this approach to binary dependent variables (Braumoeller and Carson, 2011; Xiang, 2010) survival data (Clark and Regan, 2003) and discrete ordered dependent variables (Bagozzi et al., 2015).

While the use of inflated models has accordingly become commonplace in empirical conflict studies, their applicability to the full range of dependent variables in international relations research has been limited thus far. Notably, while inflated models are now widely available for limited dependent variables that follow binary, count, ordered or duration data-generating processes, no such models have been developed and applied to "multinomial choice"-dependent variables, which seek to capture discrete *unordered* outcomes with more than two categories. This is a major limitation to empirical international relations research, as scholars in this area frequently examine dependent variables that can be best represented as unordered, polytomous choices. Consider, for instance, empirical studies of third-party war joining (e.g. Aydin, 2008), where researchers are interested in explaining whether a third country (country C) chooses to support country A, country B or neither side (i.e. status quo), during an ongoing militarized dispute between countries A and B. Similar to the "irrelevant dyads" logic presented above, it is unlikely that every country in the world has the opportunity to join an ongoing conflict between, say, India and Pakistan. Yet scholars of war joining typically include every available country-year as a possible war joiner in their endeavors. In doing so, a preponderance of irrelevant countries is likely to inflate the baseline status quo category of such dependent variables—biasing one's coefficient estimates in the exact same manner as described above.

To address these concerns, the present study considers the applicability of a newly developed baseline-inflated multinomial logit (BIMNL) model (Bagozzi and Marchetti, 2013) to outcomes of discrete unordered choice in international relations research. Building upon extant zero-inflated models for discrete binary and ordered outcomes (Beger et al., 2011;

Bagozzi et al., 2015; Harris and Zhao, 2007), the BIMNL model accounts for baseline category inflation by simultaneously estimating (a) an observation's likelihood of being baseline inflated (via a binary logit equation) and (b) an observation's probability of selecting each outcome of nominal choice—conditional on not being baseline inflated—via a multinomial logit (MNL) equation. By estimating these two processes within a single system, the BIMNL model is able to recover unbiased covariate estimates for one's polytomous dependent variable of interest (Bagozzi and Marchetti, 2013). I accordingly evaluate the effectiveness of the BIMNL for international relations research via replications of two existing MNL analyses. The first corresponds to dyads' annual decisions with respect to ongoing international territorial disputes (Huth and Allee, 2002), while the second examines the effects of political, geographic and economic factors on countries' likelihoods of experiencing various types of civil war and peace (Buhaug, 2006). In brief, I find that the BIMNL model not only exhibits generally superior model fit to the MNL model in these applications, but also yields a number of unique and theoretically intriguing findings.

Theoretical motivation

International relations scholars have long been interested in outcomes of discrete polytomous choice. As alluded to above, several past studies in this area have addressed whether (and why) a country chooses to join side A, to join side B or to sit out a given international dispute (Aydin, 2008), military alliance (Warren, 2010) or trade agreement (Hafner-Burton, 2005). Likewise, researchers interested in questions of international political economy have examined countries' polytomous exchange rate regime choices in similar manners (Bernhard and Leblang, 1999), as well as individuals' receipts of various remittance flow types, including no receipts (Adams and Cuecuecha, 2010). Conflict onset and related conflict processes have also been operationalized as unordered polytomous outcomes in scholarly efforts to assess the likelihood that a country, or pair of countries, experiences particular subtypes of interstate (or intrastate) conflict, relative to peace (Bennett and Stam, 2004; Buhaug, 2006; Buhaug and Gleditsch, 2008). Comparably, Singh and Way (2004) examine states' yearly decisions to adopt various levels of nuclear weapons programs (i.e. explore, pursue or acquire) relative to no pursuit of such weapons. Each of the above examples envisions a dependent variable with a small set of discrete, unordered outcomes or choices. Accordingly, quantitative studies of such variables have come to favor the use of nominal choice models—most frequently the MNL—to estimate the effects of covariates on an observation's probabilities of choosing each option over the others.

A second commonality shared by the above examples, however, is the presence of a "status quo" or "peace" outcome category, representing instances wherein a government or rebel group favored doing nothing, or abstaining, rather than choosing either option A or option B. While the inclusion of these status quo responses is necessary to ensure an unbiased sample of one's dependent variable, they are usually of less interest to the researcher than are the other "active" political choices, and accordingly, the vast majority of multinomial choice models in international relations treat these status quo outcomes as the baseline (i.e. reference) category in MNL estimation and interpretation. I argue below, however, that international relations scholars often risk inflating this MNL baseline category with an excess number of status quo observations that are extremely unlikely to take on any other choice outcomes under prevailing circumstances. In such instances, one's MNL estimates will be biased downward, as a significant portion of status quo cases will be impervious to the effects of covariates on transition probabilities from status quo (or peace) to other choice categories. Furthermore, when covariates have contrasting effects on the probabilities of both inflation and active multinomial choice, MNL coefficients may become further biased in indeterminate directions (Bagozzi and Marchetti, 2013).

To illustrate these concerns, consider an example from the literature on intrastate war. Recently, civil war scholars have come to argue that civil war onset is best studied in a disaggregated, multinomial fashion, wherein country-years are nominally categorized based upon whether they experience (a) territorial (i.e. separatist) rebellion, (b) governmental (i.e. overthrow-oriented) rebellion or (c) peace. Owing to the rarity of civil war, however, this approach typically entails that scholars collect global country-year samples wherein the resultant nominal "peace" outcomes outweigh rebellion outcomes by a factor as high as 70-to-1. For example, Buhaug (2006: 699) conducts such an analysis for the years 1945–1999 and reports that "there were 80 outbreaks of territorial conflict and 123 onsets of governmental conflict among the 5,411 valid country-year observations", yielding a sample composed of 97.3% peace-years. While some subset of these peace-years assuredly corresponds to conflict-prone country-years, the vast majority are likely to encompass countries such as New Zealand or Denmark, whose likelihood of experiencing a major sustained rebellion of any sort is near zero for all years under observation. Because these harmonious country-years are effectively immune to the conflict-inducing effects of short-term political, economic and climatic shocks on rebellion, the inclusion of such countries within the baseline "peace" category of MNL models of conflict onset is likely to attenuate the potentially significant effects that these temporary shocks may have upon conflict-prone countries. Moreover, by ignoring baseline inflation of this sort, scholars may similarly misattribute the inflation-inducing effects of structural factors—such as GDP per capita, which at its extremes has been suggested to induce high levels of zero-inflation bias vis-à-vis civil conflict (Bagozzi et al., 2015)—as having direct effects on civil conflict onset. This is a comparable threat to inference, in that such misattributions could lead to inaccurate causal conclusions. For example, erroneously identifying a negative direct effect of GDP per capita on civil conflict, when in fact that effect was arising solely through zero inflation, could lead researchers to conclude that development reduces societal actors' (grievance-based) *incentives* to initiate conflict against governments, whereas in reality low development simply constrains these actors' material *abilities* to rebel through, for example, extreme deprivation.

These claims are supported by extant research in the arena of international conflict, wherein zero inflation is regarded as a pervasive problem. In particular, temporal aggregations of militarized conflict—whether measured at the dyad, country or sub-national level—have been frequently found to be "inflated" with structural zeroes (Clark and Regan, 2003; Pevehouse, 2004). As mentioned above, these zeroes most commonly represent peace-observations that have a near-zero probability of experiencing conflict under contemporary circumstances. Accordingly, treating these cases as "peace-zeroes" within a statistical model of conflict can lead to biased inferences because such cases effectively have zero probability of *ever* experiencing an event of interest (Clark and Regan, 2003; Lemke and Reed, 2001; Xiang, 2010). On the other hand, truncating all potential structural (peace-year) zeroes from one's sample excludes a significant proportion of relevant-conflict observations (Bennett and Stam, 2004: 61) and produces selection bias (Lemke and Reed, 2001; Xiang, 2010). As an alternative to these approaches, scholars have now widely recognized that—by (a) including *all* observations in one's analysis and (b) then accounting for the likelihood of zero-inflation among peace observations probabilistically—one can address the challenges created by structural zeroes in an unbiased fashion (e.g. Benini and Moulton, 2004; Clark and Regan, 2003; Pevehouse, 2004; Xiang, 2010). In essence, this approach allows one to use ex-ante observable and theoretically informed covariates to account for the probability that a given zero observation is structural, and to then probabilistically discount these structural zeroes' leverage within one's primary analysis, without dropping these observations entirely.

The zero-inflated conflict framework described above has also proven useful in studies of intrastate political violence (e.g. Hegre et al., 2009; Holmes et al., 2007; Piazza, 2011). For example, in a department-month study of human rights violations committed by the Revolutionary Armed Forces of Colombia (FARC), Holmes et al. (2007) found that there were many Colombian department-months in their sample wherein the FARC was wholly inactive. The authors accordingly accounted for these structural zero-observations with a zero-inflated count model, since the FARC was probably incapable of committing *any* number of human rights violations greater than zero in departments where it was not active, and found that doing so yielded valuable theoretical and statistical insights into the underlying dynamics of civil conflict onset and intensity. At the country-year level, advanced industrialized polities have similarly been shown to engender a non-negligible quantity of structural peace-zeroes within ordinal variables of government repression and civil war (Bagozzi et al., 2015). As above, omitting all such advanced-industrialized countries from one's analysis is likely to produce selection bias and exclude a non-negligible number of actual (and potential) instances of civil conflict. Indeed, even within advanced industrial democracies, minority groups occasionally rebel against the state and homegrown terrorist attacks can occur. For those interested in producing accurate and comprehensive assessments of civil war, these are very costly cases to ignore. Thus, when faced with the potential of excess zeroes in civil conflict applications, research suggests that analysts should again (a) include all zero observations in the model and (b) account for any resultant zero-inflation econometrically.

If inflated status quo (e.g. "peace") observations arise within the multinomial choice setting in a similar fashion—and at a comparable rate—to that which has been observed in the dependent count and ordinal variable settings, then quantitative international relations studies of nominal outcomes will face the same (zero) inflation problems that past conflict researchers have widely identified. As a result, standard MNL approaches to the analysis of nominal international relations outcomes will be prone to biased findings. This will be the case no matter whether one's status quo category is treated as the MNL's baseline (i.e. reference) category or as one of the primary choice outcomes, although for ease of exposition I focus on the baseline inflation case here. To assess my claims, I explore a recently developed BIMNL model (Bagozzi and Marchetti, 2013), which—in a comparable fashion to zero-inflated count models—accounts for inflated cases via the estimation of a "first stage" logit equation. In essence, the logit equation to the BIMNL model estimates each observation's latent propensity for inflation, and then conditions the BIMNL model's subsequent MNL-stage coefficient estimates upon these estimated inflation probabilities. Given the benefits of this general framework to past international relations research, the following sections elaborate upon the BIMNL model in more detail and then evaluate my baseline inflation expectations via replications of two published studies of nominal international relations processes.

The BIMNL model

This section summarizes Bagozzi and Marchetti's (2013) BIMNL model, which was originally conceived of as a model for the heterogeneous choice processes underlying candidate preference in American politics survey data. Akin to the zero-inflated models mentioned above, the BIMNL estimator combines two latent equations: an "inflation stage" logit equation for a binary non-inflation indicator s_i and an "outcome stage" MNL equation for $\tilde{Y}_i$ (with discrete *unordered* values of $0, 1, \ldots J$), given $i \in \{1, 2, \ldots, N\}$ observations (e.g. countries). In the context of a country-year civil war dataset, for example, the status quo responses in the inflated regime ($s_i = 0$) would include peaceful countries that effectively have zero probability of experiencing civil war, while responses in regime 1 ($s_i = 1$) include occasionally peaceful countries whose probability of transitioning to a civil war outcome—given the right temporary conditions—is *not* zero. Neither $\tilde{Y}_i$ nor s_i are observable in terms of the observed baseline outcomes. However, they are observed by the criterion $Y_i = \tilde{Y}_i \times s_i$, implying that the (baseline) outcome $Y_i = 0$ can occur when $s_i = 0$ or when $s_i = 1$ and $\tilde{Y}_i = 0$. The baseline-inflated MNL distribution therefore arises as a mixture of a degenerate distribution in the baseline category and the assumed distribution of polytomous variable $\tilde{Y}_i$:

$$\Pr(Y_i) = \left\{ \begin{array}{c} \Pr(s_i = 0|z_i) + \Pr(s_i = 1|z_i)\Pr(\tilde{Y}_i = 0|x_i, s_i = 1) \text{ for } j = 0 \\ \Pr(s_i = 1|z_i)\Pr\left(\tilde{Y}_i = j|x_i, s_i = 1\right) \text{ for } j = 1, 2, \ldots J \end{array} \right\} \tag{1}$$

where $\boldsymbol{z}_i$ and $\boldsymbol{x}_i$ are inflation and outcome stage covariates, respectively. Assuming identically and independently distributed errors in these two stages, the BIMNL model is defined as:

$$\Pr(Y_i) = \left\{ \begin{array}{c} \Pr(Y_i = 0|\boldsymbol{x}_i, \boldsymbol{z}_i) = [1 - \Lambda(\boldsymbol{z'}_i\gamma)] + \left(\frac{\Lambda(z'_i\gamma)}{1 + \sum_{j=1}^{J} e^{x'_i\beta_j}}\right) \text{for } j = 0 \\ \Pr(Y_i = j|\boldsymbol{x}_i, \boldsymbol{z}_i) = \left(\frac{\Lambda(z'_i\gamma)e^{x'_i\beta_j}}{1 + \sum_{j=1}^{J} e^{x'_i\beta_j}}\right) \text{for } j = 1, 2, \ldots J \end{array} \right\} \tag{2}$$

where $\Lambda(\cdot)$ is the logistic c.d.f. Equation (3) provides the full probabilities of the BIMNL model, wherein the probability of observing outcome-choices is now conditioned upon the probability of an observation being assigned a baseline-inflated value in the multinomial d.g.p. As a result, when an unordered dependent variable is baseline-inflated, the BIMNL model provides more accurate estimates relative to a standard MNL model, in that the BIMNL estimates exhibit both less bias and greater coverage probability (Bagozzi and Marchetti, 2013).

The BIMNL's log-likelihood is derived in Bagozzi and Marchetti (2013) and can be consistently and efficiently estimated using maximum likelihood—yielding asymptotically normally distributed maximum likelihood estimates. Because the BIMNL's estimation structure is equivalent to that of extant zero-inflated models, parameter identification for this estimator is achievable even without meeting an exclusion restriction—although such restrictions are still recommended given the BIMNL model's sensitivities to misspecification and weak identification. To this end, a number of model fit statistics allow researchers to accurately test between the MNL and BIMNL models. In particular, Bagozzi and Marchetti (2013) demonstrate via Monte Carlo simulations that generalized likelihood ratio (LR) tests, Akaike information criterion (AIC), Bayesian information criterion (BIC), a variant of the Vuong test for non-nested models (Vuong, 1989) and proportion reduction in error (PRE) statistics each offer some merit in properly distinguishing between the BIMNL and MNL models. However, because the authors find disagreement between these statistics in even the most ideal of settings, they also recommended that researchers employ a combination of the test statistics mentioned above in actual applications, and that one favor the AIC/BIC and LR tests over the Vuong test and PRE.

Building upon equation (2), as well as the quantities of interest derived from similar inflated models in past research (Bagozzi and Mukherjee, 2012; Bagozzi et al., 2015; Harris and Zhao, 2007), a number of BIMNL predicted probabilities—and first differences—are available to researchers interested in illustrating substantive covariate effects. I derive the full set of probabilities—to the best of my knowledge for the first time—in the Appendix. The present section summarizes each quantity, and I further illustrate many of these predicted probabilities in my applications further below. To begin, an initial set of BIMNL quantities of interest are "inflation probabilities", which correspond to the effect of one's inflation stage covariates (and changes thereof) on the probability that an observation is inflated (vs non-inflated), or in other words, is structurally predisposed from registering any non-baseline outcome-choice. Formulas for these probabilities are presented in equations (A.2) and (A.3) of the Appendix, and can be implemented in an out-of-sample context using parametric bootstraps while holding one's inflation stage covariates to their means or medians.[1]

Researchers are often interested in assessing the effects of covariates on the BIMNL's estimated multinomial choice probabilities. In such cases, two types of BIMNL probabilities are relevant, depending on one's theoretical arguments. The first are "global choice probabilities", which communicate the full inflation *and* outcome-stage effects of a given covariate change or value—again preferably while holding other inflation and outcome stage covariates to their means or medians. In evaluating these combined (inflation and outcome stage) effects, the "global probabilities" defined in equations (A.5)–(A.10) thereby illustrate the aggregate effect of a given covariate on a "real world" observation by simultaneously accounting for that observation's inflation propensity *and* its corresponding outcome stage covariate's *inflation-conditioned* effects. Most interesting in this regard—as illustrated in the ZIOP(C) setting by Bagozzi et al. (2015)—may be instances where a variable of interest is included in *both* the inflation and outcome stage of the BIMNL model, in which case one's "global choice probabilities" can illustrate novel non-monotonic effects. Lastly, a related but distinct set of quantities of interest are "non-inflated choice probabilities", which communicate the effects of one's BIMNL outcome stage covariates on a (hypothetical) fully non-inflated observation. These "non-inflated choice probabilities" (equations A.11–A.14) most directly correspond to the coefficient estimates reported in the outcome stage of the BIMNL model, and are often the primary theoretical quantity of interest for the researcher.[2]

Applications

This section first compares the BIMNL and MNL models via a replication of Huth and Allee (2002). In their article, Huth and Allee examine the effects of domestic political accountability, namely democracy, on countries' decisions with respect to long-standing international territorial disputes. Specifically, the authors use a directed dyad-year dataset encompassing pairs of countries engaged in territorial disputes (1919–1995) to examine whether a potential territorial status quo challenger state chooses to (a) propose talks with an opposing (i.e. target) state over a disputed territory, (b) initiate a militarized interstate dispute (MID) against the target state or (c) do nothing (i.e. maintain the status quo) for a given year.[3] Huth and Allee analyze this three-category dependent variable with an MNL model, primarily treating territorial status quo as the baseline category. I focus below on the authors' "Political Accountability Model—Comparing Differences across Regimes" specification (Huth and Allee, 2002; Table 1), although I evaluate alternate specifications in the Online Appendix.

Building upon the democratic peace literature, Huth and Allee expect (and find) that domestic political accountability—measured in this case by the challenger's *democracy level*, and its interaction with *ethnic ties*, and the presence of a *recent stalemate* in negotiations—will have a negative effect on the likelihood of a challenger country initiating a territorial MID and a positive effect on the likelihood of a challenger initiating talks over the disputed territory.[4] In their analysis, the authors take care to control for each dyad's broader international strategic environment by including measures of a dyad's *common security ties* and relative *military balance*, of the *strategic value* of the territory, of whether the target or challenger were involved in other disputes, and for the *months since last challenge*. Huth and Allee generally find that these controls perform as expected, in that relative to status quo choices, *common security ties* reduce the likelihood of MIDs, *strategic value* raises the likelihood of either talks or MIDs, and increases in *military balance* (in favor of the challenger) increase the likelihood of that challenger initiating a MID.

While Huth and Allee (2002: 756) correctly note that the nature of their "territorial dispute dyad" sample helps to minimize the presence of "irrelevant dyads", extant theory nevertheless suggests that the authors' status quo observations—which encompass roughly 67% of all cases—may still include a large number of inflated dyad-years, wherein a challenger state was disproportionately incapable of initiating *any* talks or MIDs over a given disputed territory. In this respect, a number of factors may limit potential territorial dispute challengers'*opportunities* to initiate any form of territorial challenge, even if these states have the incentives or *willingness* to do so.[5] In some cases, the status quo can become a *de facto* settlement for both sides owing to increasing norms of acceptance, even while territorial disputes remain on the books in a *de jure* sense. For example, the USA and Canada currently have multiple territorial disputes (Gray, 1997; Griffiths, 2010; O'Connor, 2012), some of which are included in Huth and Allee's sample and analysis. Given the deep-rooted political and geographic stability of these two nations' shared borders, such disputes presumably would be highly constrained from rising to the level of either a high-level political challenge or an outright MID under current conditions. Severe power asymmetries may have similar inflation-inducing effects, as relative deficiencies in the material capabilities of (and hence opportunities for) credibly challenging disputed territories have likewise been identified as severely constraining states' abilities to challenge disputed territorial claims.[6] Lastly, geographic interaction opportunity (e.g. contiguity) has been widely shown to be *both* a strong predictor of zero inflation within MIDs (Bagozzi et al., 2015; Xiang, 2010) *and* a significant constraint upon states' abilities to initiate territorial challenges (Huth, 1998; Senese, 2005)—implying that contiguity may have similar effects within Huth and Allee's sample. Moreover, while the authors' "territorial dispute" sample probably reduces these geographic effects to some degree, the fact that approximately 35% of the cases in their sample are noncontiguous suggests that geographic distance may indeed still leave some challenger countries disproportionately unable to address disputed territories in any manner.

Table 1. BIMNL and MNL models of challenger decisions to challenge the status quo

	MNL		BIMNL		
	Talks vs no action	Force vs no action	Talks vs no action	Force vs no action	Pr(Non-inflation)
Constant	–0.743*** (0.082)	–3.209*** (0.179)	4.324*** (0.927)	1.692* (0.931)	–0.733*** (0.095)
Challenger democracy level	0.018*** (0.006)	–0.055*** (0.013)	–0.038 (0.050)	–0.113** (0.051)	0.013* (0.008)
Challenger democracy × stalemate	–0.003 (0.010)	–0.014 (0.019)	0.067 (0.043)	0.051 (0.046)	–0.019 (0.013)
Recent stalemate	0.449*** (0.123)	0.332 (0.205)	–0.766 (0.668)	–0.702 (0.681)	0.616*** (0.163)
Challenger democracy × ethnic ties	0.015* (0.008)	0.050*** (0.017)	0.004 (0.045)	0.041 (0.047)	0.036*** (0.011)
Ethnic ties	0.140 (0.098)	0.253 (0.172)	–1.458* (0.838)	–1.311 (0.846)	0.274** (0.136)
Common security ties	–0.052 (0.064)	–0.423*** (0.126)	0.019 (0.324)	–0.338 (0.342)	–0.186** (0.084)
Strategic value	0.199*** (0.068)	0.381*** (0.120)	1.632** (0.733)	1.632** (0.733)	1.775** (0.740)
Challenger in other dispute	–0.101 (0.075)	0.492*** (0.125)	0.011 (0.284)	0.651** (0.298)	
Target in other dispute	–0.016 (0.068)	0.340*** (0.124)	0.717 (0.606)	1.018* (0.611)	
Military balance	–0.032 (0.116)	1.438*** (0.210)	–3.969*** (0.909)	–2.406** (0.919)	0.829*** (0.278)
Months since last challenge	–0.006*** (0.001)	–0.016*** (0.002)	–0.000 (0.002)	–0.007*** (0.002)	–0.007*** (0.001)
Contiguity					–0.039 (0.074)
Major power					0.248** (0.074)
PRE	0.3%		2.0%		
AIC	9646.8		9629.6		
BIC	9809.7		9867.0		

Notes: $N = 6544$. Standard errors in parentheses. ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$. Vuong test = -2.515 ($p < 0.05$). LR test = 39.34 ($p < 0.01$).

If any of the above inflated status quo cases are present, then one's ability to estimate unbiased and consistent variable effects on a challenger's decision to initiate territorial talks or MIDs will be severely undermined. To evaluate these claims, I replicate the analysis reported in Huth and Allee (2002; Table 1) using MNL and BIMNL models.[7] I keep each model's outcome stage equivalent to that reported by Huth and Allee (2002), and then include a number of theoretically informed covariates in my BIMNL inflation stage. Specifically I add Huth and Allee's *challenger democracy level* variable, and all related interaction components, to the inflation stage under the assumption that this variable and its interactive effects will help to capture the institutional and norm-based constraints that may increasingly preclude countries from realistically challenging the status quo. Second, and given (a) the opportunity-based territorial constraints discussed above and (b) extant zero-inflation findings for MIDs, I then also include *common security ties*, *contiguity* and *major power* in my inflation stage, as the first of these three covariates has been found to make dyads significantly less able to have international disputes (Bagozzi et al., 2015), whereas the latter two have been argued to make dyads more capable of interacting and initiating disputes (Bagozzi et al., 2015; Xiang, 2010), including territorial disputes (Gibler, 2007; Huth, 1998; Senese, 2005). Finally, I add *military balance* to my inflation stage in order to account for the aforementioned potential that relatively weak challengers may be simply less able to reasonably challenge the status quo, and then complete the inflation stage specification with *months since last challenge*, which is intended to control for both temporal dependence and any increasing norms of status quo acceptance that may arise between the target and challenger.

These MNL and BIMNL results are reported in Table 1 and offer a number of insights. I first examine my inflation stage results, where the dependent variable is the probability that an observation is *non-inflated*. Doing so, I find that *common security ties* make challengers significantly less likely to be able to challenge the status quo via any means, as does *months since last challenge*. Both findings are consistent with the opportunity and norm-based expectations outlined above. Second, as *military balance* (in favor of the challenger) and *major power* increase, a challenger becomes more able to challenge the status quo in some manner (i.e. to be non-inflated), again in line with expectations, although *contiguity* surprisingly has no significant effect in this regard. Finally, it is difficult to fully interpret the democracy-interactions' inflation-stage coefficient estimates in Table 1—and more attention will be given to components of these in the predicted probability calculations below—although one can at least note in Table 1 that anocracies with *recent stalemates* appear significantly more likely to be non-inflated, and to be able to challenge the status quo via some means, as do democratic challengers to a degree. Note however that the latter finding is only significant at the $p<0.10$ level, and is conditional on *recent stalemates* and *ethnic ties* each being equal to zero. Hence, altogether, the inflation stage results are generally consistent with my theoretical expectations, and I withhold further interpretation of the implications of these inflation stage effects until extracting the combined "global choice probabilities" further below.

Before doing so, I briefly compare the BIMNL and MNL outcome stage coefficient estimates from Table 1, where importantly the BIMNL outcome stage estimates now report the effects of one's covariates on a fully non-inflated (i.e. status quo-challenge capable) observation. Beginning with the democracy interactions, I find that *challenger democracy levels* no longer exhibit a significant positive effect on territorial talks, once baseline inflation is accounted for, although this variable continues to exert a significant negative effect on the use of force. Because *challenger democracy levels* are interacted with dichotomous covariates, these findings apply to the case where *status quo* and *ethnic ties* are set to zero. With *challenger democracy* set to zero (i.e. for anocracies), *recent stalemates* intuitively no longer have a significant positive effect on the likelihood of talks, whereas *ethnic ties* now have a negative and significant direct effect on talks. This latter finding accordingly implies that, relative to opting for the status quo, anocratic challenger countries that are able to initiate *some* challenge to the status quo are more likely to initiate territorial talks (relative to maintaining the status quo) when ethnic co-nationals inhabit the disputed territory. Many of additional outcome stage covariates are comparable across the BIMNL and MNL models in both their signs and general levels of statistical significance, including target/challenger involvement in other disputes and *strategic value*, and do not warrant further discussion here. On the other hand, *military balance* is now a significant negative predictor of both outcomes in the BIMNL model. This implies that challengers with superior military capabilities become less likely to initiate either form of challenge, relative to maintaining the status quo, once they are *able* to initiate a challenge. To further explore these findings, I examine the substantive significance of two key covariates of interest—*military balance* and the *strategic value* of a territory—via predicted probabilities, and changes thereof, immediately below.

To derive these predicted probabilities, I use "global choice probabilities" for the BIMNL model, which are defined in equations (A.5)–(A.10) of the Appendix. Theory should always dictate one's choice of reported probabilities, and given that the authors' sample is already constrained to their primary cases of interest (i.e. territorial dispute dyads), "global choice probabilities" are most relevant to the application at hand. Indeed, and as mentioned above, these probabilities communicate the *net effects* of one's covariates when taking into account one's inflation *and* outcome stage processes, and therefore speak directly to the effects of variable(s) across one's entire sample. Specifically using equation (A.6) to evaluate the effects of *military ratio* on the predicted probability of talks and MIDs, I derive my "global choice probabilities" for both the talks and MID outcomes when simultaneously iterating across the range of *military ratio* within the inflation *and* outcome stages of the BIMNL model. I do so while holding all other variables to their means (for continuous variables) or medians (binary and ordinal variables). Confidence intervals were obtained using parametric bootstraps ($M = 1000$) in a manner comparable to that described in Krinsky and Robb (1986).

Plotting these quantities in Figure 1, I first find that an increase in *military ratio* (and hence the challenger's relative military superiority) monotonically increases the likelihood of territorial conflict in both the MNL and BIMNL contexts. This positive relationship is consistent with past findings (Huth, 1996; Huth and Allee, 2002), although it is less nuanced than one might expect, given that bargaining theory suggests a non-monotonic relationship between military power and interstate conflict wherein, at high levels of military preponderance, conflict should again become less likely (Reed, 2003). Note however that—in contrast to the MNL conflict-predicted probabilities—the BIMNL conflict probabilities *do* appear to be decelerating at high levels of challenger military preponderance, which suggests that the BIMNL model results are (slightly) more consistent with these bargaining theory expectations. Further, traditional conflict bargaining dynamics may not perfectly hold in cases of disputed territory, wherein very strong challengers probably have both the military capabilities to easily seize disputed territory and the incentives to favor this approach relative to talks. Indeed, as others have noted, militarily preponderant territorial dispute challengers do at times favor military action over talks, with notable cases being Iraq's invasion of Kuwait and the Sino-Indian war (Huth, 1996). At the same time, the MNL and BIMNL models yield notably different predicted probabilities for the likelihood of territorial talks: while the MNL

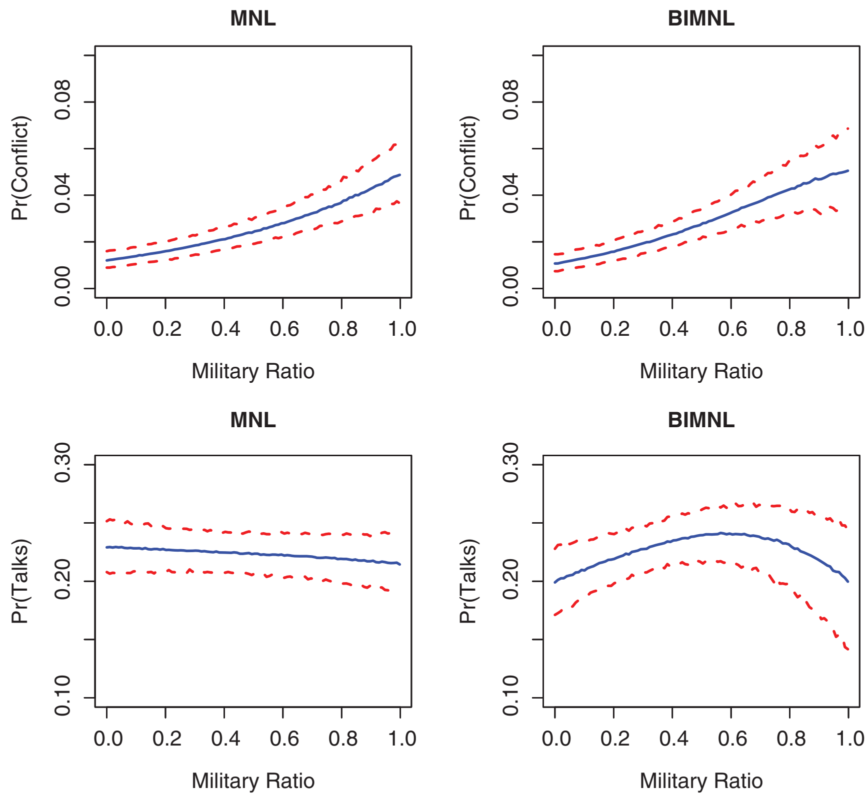

Figure 1. Predicted probabilities of challenger decisions to challenge the status quo across range of military ratio.

model reports the effect of *military ratio* on talks to be insignificant, the BIMNL model suggests that increases in *military ratio* from the lowest levels of relative challenger military capabilities to challenger–target military-parity increase the likelihood of talks, whereas further increases in *military ratio* then decrease the likelihood of such talks, as a challenger's increasingly superior military capabilities ostensibly lead it to favor the use of MIDs over talks. These nonmonotonic effects, which go entirely unnoticed when baseline inflation is ignored in the MNL model, are consistent with extant research in this arena (Allee and Huth, 2006; Huth, 1996; Hensel, 2001), which suggests (and finds) that dyads exhibiting military parity are more likely to pursue cooperative territorial dispute resolution tactics relative to maintaining the status quo. Explanations for this tendency posit that militarily disadvantaged challengers initially have little recourse to demand talks (and instead favor the status quo), but increase such demands as their military advantages increase, before ultimately decreasing their calls for talks relative to military action given that they increasingly pose a distinct advantage in the latter approach (Allee and Huth, 2006; Huth, 1996).

I next examine the *changes* in "global choice probabilities" that one obtains from the BIMNL and MNL models when the *strategic value* of a territory is changed from zero to one. In this case, because *strategic value* is not included in the inflation stage of the BIMNL

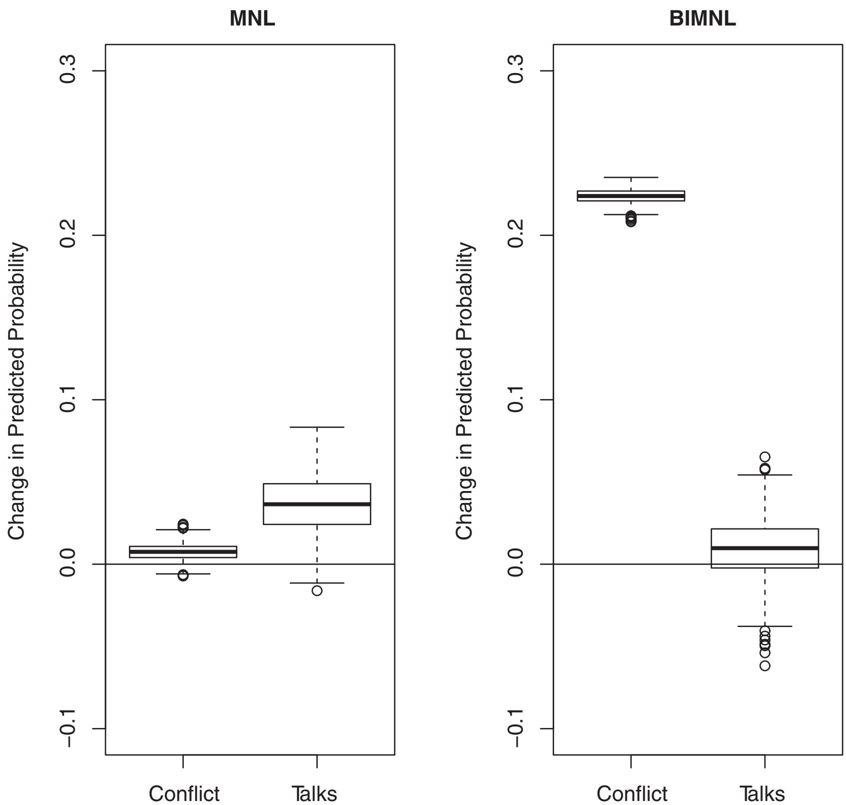

Figure 2. Change in predicted probabilities of challenger decisions to challenge the status quo given a zero-to-one change in strategic value.

model, *strategic value* is only modified within the outcome equation, and all inflation stage (and remaining outcome stage) covariates are held to mean or median values. Presenting these results via boxplots in Figure 2, I find that, in accounting for baseline inflation, the BIMNL estimated-effect of a zero-to-one change in *strategic value* on the probability of territorial conflict is substantially higher than is the case for the MNL model. Specifically, a zero-to-one change in *strategic value* is expected to increase the probability of a MID in the MNL model by an insignificant 1%, whereas in the BIMNL model this effect is significant and substantial, with an average estimated increase of 22% in the predicted probability of conflict. This latter result, and its magnitude, is more consistent with past findings of *strategic value*'s strong positive effects on challengers' *willingness* to dispute a given territory (Huth, 1998: 73), as well as arguments thereof (Huth, 1998), especially when strategic value is interpreted as salience (e.g. Hensel, 2001). Thus, by not accounting for inflation, the MNL model exhibits several instances in which the substantive magnitude of statistically significant covariates is severely underestimated.

Lastly, one can also compare the MNL and BIMNL models via the model fit statistics reported in Table 1 to gain a better sense of which model more appropriately fits the data.

Doing so, I find that a majority of the model fit statistics mentioned earlier favor the BIMNL model over the MNL model. For instance, the Vuong test strongly favors the BIMNL model over the MNL ($p<0.05$) as does a PRE statistic. Similarly, the AIC and LR test—two of the three preferred model fit statistics (Bagozzi and Marchetti, 2013)—each favor the BIMNL model over the MNL model, although the BIC favors the MNL model. Taken together, these model fit statistics moderately suggest that the BIMNL specification presented above provides a superior fit to the data compared with a comparable MNL model, implying that my accounting for baseline inflation in this case was probably warranted. Additional BIMNL model specifications for this replication in Tables A.1 and A.2 of the Online Appendix demonstrate that my BIMNL model and the results discussed above are fairly robust to alternative model specifications.

My second replication examines a recent article by Buhaug (2006) that sought to assess the effects of a number of political–economic covariates—such as political institutions, economic development and oil dependence—upon a country's likelihood of experiencing peace, territorial rebellion or governmental rebellion in a given year. Buhaug's dependent conflict variable accordingly takes on three discrete, unordered choices—territorial civil war onset, governmental civil war onset and peace—and the author constructs a global country-year tally of these choices for the years 1946–1999. Consequently, Buhaug's sample exhibits an extreme number of peace country-year cases, encompassing roughly 97% of all observations. I plot the full distribution of this dependent variable in the Online Appendix. As argued above and in extant research (Bagozzi et al., 2015), a majority of the domestic-peace observations in these contexts probably correspond to harmonious country-years—such as contemporary Denmark or Sweden—that cannot reasonably experience any degree of full-scale rebellion. By including these observations within the baseline "peace" category of a multinomial logit model, as Buhaug (2006) does, one potentially risks under (or over) estimating the direct effects of variables such as polity and GDP per capita on conflict—as comparable variables have been shown to affect peace-inflation probabilities in previous studies of civil war (Bagozzi et al., 2015).

To evaluate this possibility, I re-estimate Buhaug's MNL specification using the BIMNL model and compare each model's results. I do so while including the covariates used in Buhaug's primary MNL model within my BIMNL model's outcome stage: *democracy*, *anocrocy*, *GDP per capita*, *oil exporter*, log *country size*, *ethnic fractionalization* and a conflict/independence *decay function*. All variable operationalizations are described in the Online Appendix; relevant independent variables are lagged by one year. Drawing on past research, I also identify and include a number of relevant covariates in the inflation stage of my BIMNL model. In particular, I follow Bagozzi et al. (2015) to include *oil exporter* and *GDP per capita* in my inflation stage, as extreme oil wealth may provide governments with the resources to suppress any level of rebellion (Basedau and Lay, 2009) while countries with very high levels GDP per capita may similarly be robust to major rebellion (Bagozzi et al., 2015). I complete my inflation stage specification with controls for regime type (*polity*) and temporal dependence (via *decay function*).

The MNL and BIMNL results appear in Table 2. Here, I begin by comparing the outcome stage estimates for these two models. In this regard, one can note that several variables have consistent effects across both the MNL and BIMNL models. For instance, the two models similarly suggest that *country size* and *ethnic fractionalization* each significantly and positively affect the likelihood of territorial civil war (relative to peace)—but have no significant effect on governmental civil war in either model. In addition, *GDP per capita*, as

Table 2. BIMNL and MNL models of rebellion

	MNL		BIMNL		
	Talks vs no action	Force vs no action	Talks vs no action	Force vs no action	Pr(Non-inflation)
Constant	–12.480*** (1.188)	–4.404*** (0.760)	–11.898*** (1.306)	–3.644*** (1.014)	
Democracy	0.896*** (0.312)	–0.133 (0.288)	0.690 (0.457)	–0.469 (0.538)	
Anocracy	0.228 (0.300)	0.568*** (0.206)	0.131 (0.391)	0.354 (0.365)	
GDP per capita	–0.167** (0.063)	–0.158*** (0.050)	–0.246*** (0.078)	–0.274*** (0.076)	0.388** (0.149)
Oil exporter	0.585* (0.321)	0.838*** (0.241)	1.643*** (0.623)	2.190*** (0.636)	–2.460*** (1.147)
Country size	0.493*** (0.087)	0.035 (0.060)	0.503*** (0.089)	0.046 (0.061)	
Ethnic Fractionalization	1.566*** (0.485)	0.539 (0.338)	1.702*** (0.504)	0.575 (0.355)	
Decay function	3.169*** (0.316)	0.890*** (0.291)	2.541*** (0.475)	0.113 (0.517)	2.661*** (0.784)
Polity					0.095 (0.065)
PRE	2.0%		1.0%		
AIC	1724.6		1723.5		
BIC	1830.2		1862.0		

Notes: $N = 5411$. Standard errors in parentheses. ***$p < 0.01$; **$p < 0.05$; **$p < 0.1$. Vuong test $= -1.47$ ($p < 0.14$). LR Test $= 11.13$ ($p < 0.05$).

Buhaug hypothesized and found, has a negative and significant effect on both territorial and governmental conflict across the BIMNL and MNL models, whereas *oil exporter* has a significant conflict-inducing effect within both models. However, these latter BIMNL results, which now represent the effect of each of these covariates among only those (non-inflated) countries that could reasonably experience conflict during these periods, are now much more substantively sizable than before.

To illustrate this point, I present the estimated effects of reasonable changes in *oil exporter* and *GDP per capita* on the out-of-sample predicted probabilities of observing each type of civil war via boxplots in Figure 3. As above, I do so while holding all other variables to their means (for continuous variables) or medians (binary and ordinal variables) and derive confidence intervals using parametric bootstraps ($m = 1000$). Of most interest in this case are the estimated effects of these variables among only those hypothetical countries that could reasonably experience civil war and I therefore compare my MNL and BIMNL models using the "non-inflated choice probability" formula in equation (A.14) of the Appendix. One can note here that, for realistic changes in *oil exporter* and *GDP per capita*, the direct BIMNL estimated effect of each of these variables on the probability of governmental rebellion among conflict-capable (i.e. non-inflated) countries is several times that of the comparable MNL-estimated effect, and in each case is statistically significant at the $p < 0.05$ level. In Figure 3a for example, the MNL model suggests that a country changing from a non-*oil exporter* to an *oil exporter* could be expected to see a 2% increased risk of governmental civil war (95% CI $= 1\% \leftrightarrow 4\%$), whereas the BIMNL model implies that, conditional on a country being capable of experiencing civil war, a comparable change in *oil exporter* would lead to a 15% increase in governmental rebellion (95% CI $= 2\% \leftrightarrow 38\%$). Hence, once "irrelevant" country-year cases are accounted for in one's analysis, a number of commonly held conflict correlates appear to have significantly larger direct effects on conflict than previously realized.

The MNL and BIMNL outcome stage results reported in Table 2 further diverge when one turns to Buhaug's political institutions variables. In the original analysis, Buhaug interestingly finds that democracies are significantly *more* likely to experience territorial conflict than autocracies. In contrast, with regard to governmental rebellion, Buhaug found that democracy had no significant effect on the likelihood of rebellion, whereas anocracy increased the likelihood of rebellion. The BIMNL reveals, however, that after accounting for countries' propensities to experience any degree of rebellion, neither set of political institutions has direct effects on a country's likelihood of experiencing civil war. Turning to the inflation stage estimates for the BIMNL model, *oil exporters* are significantly less likely to be able to experience civil war of any kind (and are thus more likely to be inflated), which is in line with theoretical expectations. *GDP per capita* is negative and significant, implying that, in contrast to Bagozzi et al. (2015), increases in *GDP per capita* make countries more likely to be able to experience some form of civil conflict. However, one should note that the non-peace outcomes in Bagozzi et al. (2015) encompassed not only civil war but also government repression, which probably entails a distinct inflation process from that examined here. Finally, while *polity* is insignificant, the *decay function* is statistically significant and positive in the BIMNL's inflation stage, suggesting that countries that have more recently experienced independence and/or civil war are more likely to be capable of experiencing civil war—perhaps because these are the very countries that are the least likely to be politically and socially consolidated.

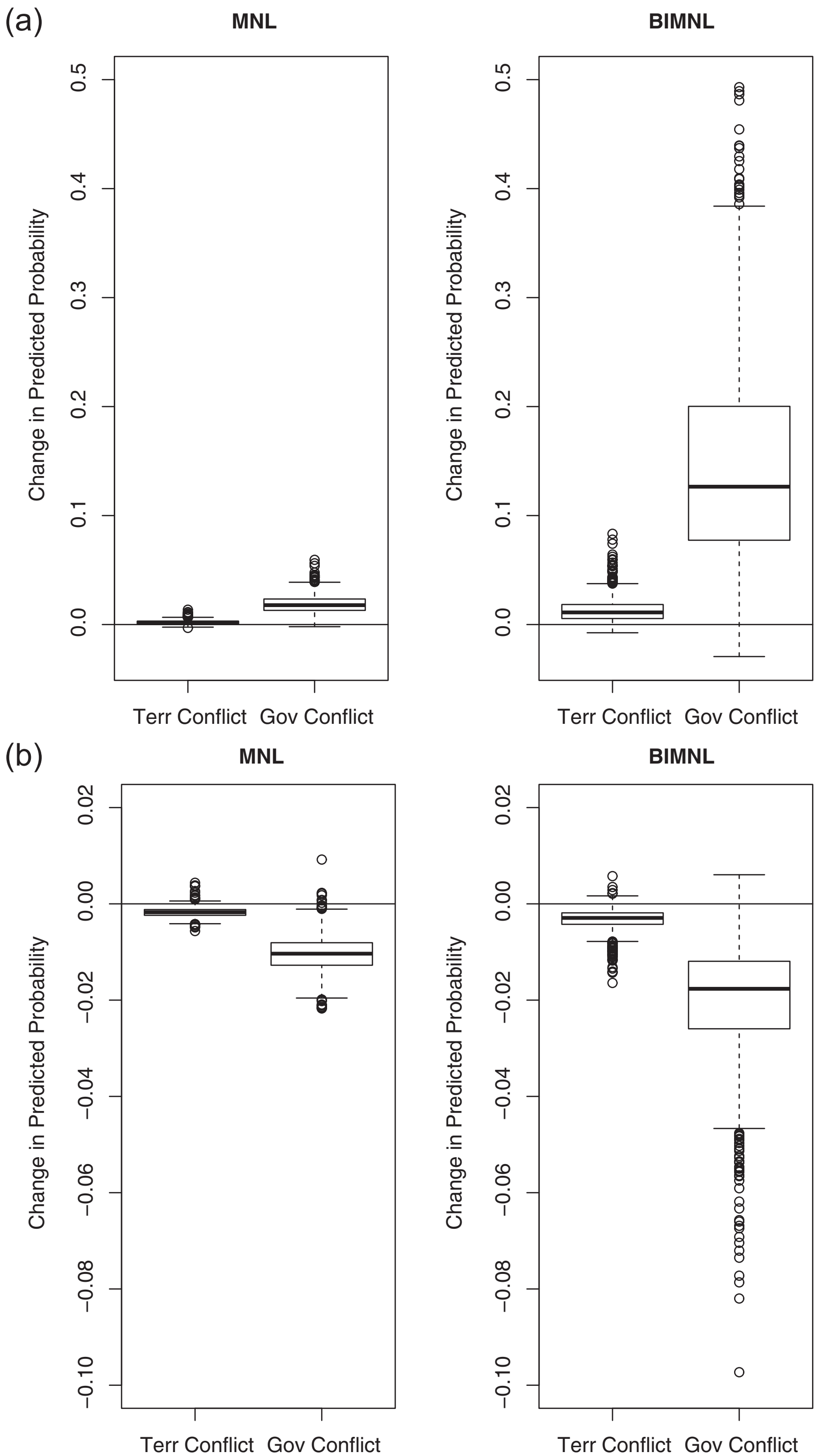

Figure 3. Changes in predicted probabilities of governmental and territorial civil war.

For this application, model fit statistics (see Table 2) are less conclusive as to whether the BIMNL model provides for a better fit than the MNL model. Two of the three preferred model fit statistics (the AIC and LR test) favor the BIMNL model over the MNL model, whereas the BIC again favors the MNL model. The Vuong test in this case is inconclusive, favoring neither model over the other, while the PRE marginally favors the MNL model. Hence, it is worth emphasizing here that the BIMNL specification presented above may not provide much of an improvement over the MNL model for this particular replication, and future applications of the BIMNL model to dependent variables of multinomial rebellion should seek to establish a broader set of structural variables for better inflation stage specification. Finally, I discuss a selection of (smaller and larger) BIMNL model specifications for this replication in Tables A.3 and A.4 of the Online Appendix, which together demonstrate that, although my model fit statistics are inconclusive, the results discussed above are nevertheless highly stable.

Conclusion

Within international relations, outcomes of discrete polytomous choice will often include heterogeneous mixtures of status quo (e.g. peace) observations. Extant zero-inflated research suggests that one subset of these status quo cases will typically correspond to observations that actively choose the status quo over all other alternatives owing to non-structural factors such as short-term political–economic shocks. On the other hand, many other status quo responses are likely to be *inflated* status quo cases that, owing to slow-moving or structural factors such as geography or GDP, are unlikely to ever register a non-status quo outcome. These latter countries (or dyads) are largely immune to the effects of non-status quo choice determinants. Ignoring this heterogeneity can severely bias the estimated effects of one's covariates on multinomial choice (Bagozzi and Marchetti, 2013). Thus, it is important for international relations researchers to account for these disparate subsets of status quo choices when "peace" or "status quo" serves as the reference category in empirical models of multinomial choice.

This paper considers a new discrete choice estimator—the BIMNL model—that is designed to allow researchers to more accurately do so. After refining this model, I demonstrate via replications of extant research that ignoring baseline category inflation within quantitative studies of international relations can significantly bias one's parameter estimates. Regarding these replications, my findings also offer new theoretical insights into the determinants of political conflict and cooperation. For instance, a reanalysis of Huth and Allee (2002) with the BIMNL model uniquely demonstrates that the effects of dyadic military ratios on countries' decisions to peacefully settle territorial disputes are significant, and non-monotonic—rising as a potential status quo challenger's military power grows to parity with a target, and then declining thereafter as military options become more enticing. Similarly, my replication of Buhaug (2006) suggests that a number of commonly studied predictors of civil conflict, namely natural resource dependence and GDP per capita, exhibit significantly larger effects on the likelihood of governmental civil war once baseline (i.e. peace) inflation has been better accounted for. As such, these findings not only offer deeper insights into the potential causal mechanisms underlying conflict processes, but may also provide an improved framework for the forecasting of multinomial conflict phenomena.

The current study can be extended in three directions. First, given the time-series cross-sectional nature of many international relations datasets, a random effects BIMNL model would further improve the utility of the BIMNL model for international relations research. Extensions of this sort have been implemented for inflated ordered dependent variable models (Brooks et al., 2012), which suggests that this extension would be both feasible and beneficial. Second, nominal international political economy variables, including those relating to country membership in international institutions, exchange rate regime choice, and remittance flow-type may be similarly prone to baseline category inflation. Analysis of these international political economy outcomes with the BIMNL model is thus an important next step, wherein a comparison of the BIMNL model to the constrained MNL model is particularly intriguing. Finally, the BIMNL model could be generalized to the multinomial probit or mixed-logit settings. This would allow researchers to account for baseline category inflation in instances where they believe that their nominal dependent variable is in violation of the Independence of Irrelevant Alternatives assumption. Moreover, a generalization of the BIMNL model to the multinomial probit could also allow one to estimate a correlated disturbance term between the two stages of one's inflated multinomial choice model. Correlated disturbances have proven useful in international relations studies of inflated ordered and binary dependent variables (Bagozzi et al., 2015; Xiang, 2010) and thus would probably be of interest to international relations scholars within the nominal setting as well.

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Notes

1. Such quantities correspond closely to the zero-inflated ordered probit with correlated disturbance's "non-harmony" probabilities presented in Bagozzi et al. (2015).
2. These are comparable to those reported in Bagozzi and Mukherjee (2012) and Bagozzi and Marchetti (2013). A similar set of quantities can be obtained by calculating odds ratios from one's BIMNL outcome stage coefficient estimates. In the interest of space, I present and discuss such odds ratios in the Online Appendix, available at: http://www.benjaminbagozzi.com/research.html
3. I plot the full distribution of this dependent variable in the Online Appendix.
4. All variables are operationalized in the Online Appendix.
5. Here, and consistent with extant zero-inflated studies (Pevehouse, 2004; Xiang, 2010), I am equating determinants of "opportunity" with determinants of zero-inflation.
6. See, for example, Gibler (2007), who in addition to referencing the role of power asymmetry in territorial conflict, further argues that "[l]arge differences in power would suggest that a border is unlikely to undergo renegotiation regardless of how the border was previously defined" (p. 520).
7. Replication data were obtained from the *Journal of Conflict Resolution* website. However, I did not have access to the original estimation commands used by the authors, and my attempts to replicate the original analysis from the available data left me with two additional observations relative to what was reported in Huth and Allee (2002) (i.e. an *N* of 6544 instead of 6542). I also chose not to

employ robust standard errors. Even so, the findings reported in my MNL replication are generally consistent with those reported in Huth and Allee (2002).

References

Adams RH Jr and Cuecuecha A (2010) Remittances, household expenditure and investment in Guatemala. *World Development* 38(11): 1626–1644.

Allee TL and Huth PK (2006) Legitimizing dispute settlement: international legal rulings as domestic political cover. *American Political Science Review* 100(2): 219–234.

Aydin A (2008) Choosing sides: Economic interdependence and interstate disputes. *Journal of Politics* 70(4): 1098–1108.

Bagozzi BE and Mukherjee B (2012) A mixture model for middle category inflation in ordered survey responses. *Political Analysis* 20(3): 369–386.

Bagozzi BE and Marchetti K (2013) Distinguishing occasional abstention from routine ambivalence in models of vote choice. Paper presented at the *APSA 2013 Annual Meeting*.

Bagozzi BE, Hill DW, Moore WH and Mukherjee B (2015) Modeling two types of peace: The zero-inflated ordered probit (ZiOP) model in conflict research. *Journal of Conflict Resolution*, in press.

Basedau M and Lay J (2009) Resource curse or rentier peace? The ambiguous effects of oil wealth and oil dependence on violent conflict. *Journal of Peace Research* 46(6): 757–776.

Beger A, DeMeritt JHR, Hwang W and Moore WH (2011) The Split Population Logit (SPopLogit): Modeling measurement bias in binary data. Working paper.

Benini AA and Moulton LH (2004) Civilian victims in an asymmetrical conflict: Operation Enduring Freedom, Afghanistan. *Journal of Peace Research* 41(4): 403–422.

Bennett DS and Stam A (2004) *The Behavioral Origins of War*. Ann Arbor, MI: University of Michigan Press.

Bernhard W and Leblang D (1999) Democratic institutions and exchange-rate commitments. *International Organization* 53(1): 71–97.

Braumoeller BF and Carson A (2011) Political irrelevance, democracy, and the limits of militarized conflict. *Journal of Conflict Resolution* 55(2): 292–320.

Brooks R, Harris MN and Spencer C (2012) Inflated ordered outcomes. *Economic Letters* 117: 683–686.

Buhaug H (2006) Relative capability and rebel objective in civil war. *Journal of Peace Research* 43(6): 691–708.

Buhaug H and Gleditsch KS (2008) Contagion or confusion? Why conflicts cluster in space. *International Studies Quarterly* 52(2): 214–233.

Clark DH and Regan PM (2003) Opportunities to fight: A statistical technique for modeling unobservable phenomena. *Journal of Conflict Resolution* 47(1): 94–115.

Gibler DM (2007) Bordering on peace: Democracy, territorial issues, and conflict. *International Studies Quarterly* 51(3): 509–532.

Gray DH (1997) Canada's unresolved maratime boundaries. *IBRU Boundary and Security Bulletin* (autumn).

Greene WH (2011) *Econometric Analysis*, 7th edn. Englewood Cliffs, NJ: Prentice-Hall.

Greene WH and Hensher DA (2010) *Modeling Ordered Choices: A Primer*. Cambridge, MA: Cambridge University Press.

Griffiths S (2010) US–Canada Arctic border dispute key to maritime riches. *BBC News*, 2 August 2010 (accessed 4 January 2014).

Hafner-Burton E (2005) Forum shopping for human rights: The transformation of preferential trade. Paper presented at the *American Political Science Association Annual Meeting*, Washington DC.

Harris MN and Zhao X (2007) A zero-inflated ordered probit model, with an application to modelling tobacco consumption. *Journal of Econometrics* 141(2): 1073–1099.

Hegre H, Ostby G and Raleigh C (2009) Property and civil war: A disaggregated study of Liberia. *Journal of Conflict Resolution* 53(4): 598–623.
Hensel PR (2001) Contentious issues and world politics: The management of territorial claims in the Americas, 1816–1992. *International Studies Quarterly* 45(1): 81–109.
Holmes JS, Gutierrez De Pineres SA and Curtina KM (2007) A subnational study of insurgency: FARC violence in the 1990s. *Studies in Conflict & Terrorism* 30(3): 249–265.
Huth PK (1996) Enduring rivalries and territorial disputes, 1950–1990. *Conflict Management and Peace Science* 15(7): 7–41.
Huth PK (1998. *Standing your Ground: Territorial Disputes and International Conflict*. Ann Arbor, MI: University of Michigan Press.
Huth PK and Allee TL (2002) Domestic political accountability and the escalation and settlement of international disputes. *Journal of Conflict Resolution* 46(6): 754–790.
Krinsky I and Robb AL (1986) On approximating the statistical properties of elasticities. *The Review of Economic and Statistics* 68(3): 715–719.
Lemke D and Reed W (2001) The relevance of politically relevant dyads. *Journal of Conflict Resolution* 45(1): 125–144.
Maoz Z and Russett BM (1993) Normative and structural causes of democratic peace, 1946–1986. *American Political Science Review* 87(3): 624–638.
O'Connor J (2012) Puffin Wars: The island paradise at centre of last Canada–U.S. land dispute. *National Post*, 27 November 2012 (accessed 4 January 2014).
Pevehouse J (2004) Interdependence theory and the measurement of international conflict. *Journal of Politics* 66(1): 247–266.
Piazza JA (2011) Poverty, minority economic discrimination, and domestic terrorism. *Journal of Peace Research* 48(3): 339–353.
Reed W (2003) Information, power, and war. *American Political Science Review* 97(4): 633–641.
Senese PD (2005) Territory, contiguity, and international conflict: Assessing a new joint explanation. *American Journal of Political Science* 49(4): 769–779.
Singh S and Way CR (2004) The correlates of nuclear proliferation: A quantitative test. *Journal of Conflict Resolution* 48(6): 859–885.
Vuong QH (1989) Likelihood ratio tests for model selection and non-nested hypotheses. *Econometrica* 57(2): 307–333.
Warren TC (2010) The geometry of security: Modeling interstate alliances as evolving networks. *Journal of Peace Research* 47(6): 697–709.
Xiang J (2010) Relevance as a latent variable in dyadic analysis of conflict. *Journal of Politics* 72(2): 484–498.

Appendix

As mentioned in the main paper, one can use the basic probability statements of the BIMNL model to generate predictions for the probability of observing each outcome category, as well as for observations' probabilities of inflation. The ideal framework for doing so matches that presented for the applications in the text, wherein one calculates out-of-sample predicted probabilities (or probability changes) after selecting and setting the values of other independent variables to a fixed set of values (e.g. means or medians). Parametric bootstraps can then be used to calculate standard errors and confidence intervals. Doing so accordingly allows one to examine (choice or inflation) predictions across the sample-range of a given independent variable, or changes in predicted probabilities given a particular change (e.g. standard deviation, or minimum to maximum) in an independent variable. To illustrate this, I first re-present the full probability statement for the BIMNL model—given a sample of

$i \in \{1, 2, \ldots N\}$ observations and a nominal dependent variable with $j = 0, 1, \ldots, J$ categories—in equation (A.1):

$$\Pr(Y_i) = \begin{cases} \Pr(Y_i = 0|\boldsymbol{x}_i, \boldsymbol{z}_i) = [1 - \Lambda(\boldsymbol{z}'_i\gamma)] + \left(\frac{\Lambda(\boldsymbol{z}'_i\gamma)}{1 + \sum_{j=1}^{J} e^{\boldsymbol{x}'_i\beta_j}}\right) \text{for } j = 0 \\ \Pr(Y_i = j|\boldsymbol{x}_i, \boldsymbol{z}_i) = \left(\frac{\Lambda(\boldsymbol{z}'_i\gamma)e^{\boldsymbol{x}'_i\beta_j}}{1 + \sum_{j=1}^{J} e^{\boldsymbol{x}'_i\beta_j}}\right) \text{for } j = 1, 2, \ldots J \end{cases} \tag{A.1}$$

where $\Lambda(\cdot)$ is the logistic c.d.f., $\boldsymbol{z}'_i$ corresponds to one's inflation stage covariates with corresponding coefficient estimates γ, and $\boldsymbol{x}'_i$ corresponds to one's outcome stage covariates, with coefficient estimates β.

The first set of predicted probabilities and probability changes of interest for the BIMNL model corresponds to the probability of inflation (i.e. $\Pr(s = 0)$), and changes therein. The former of these quantities, providing the probability of inflation given specified values of one's inflation stage covariates ($\boldsymbol{z}$), can be defined as:

$$\Pr(s = 0) = 1 - \frac{\exp(\boldsymbol{z}\hat{\gamma})}{1 + \exp(\boldsymbol{z}\hat{\gamma})} = 1 - \Lambda(\boldsymbol{z}\hat{\gamma}) \tag{A.2}$$

while the latter quantity—that is, the change in $\Pr(s = 0)$ associated with a specified change in the values of the covariate vector $\boldsymbol{z}$ from $\boldsymbol{z_A}$ to $\boldsymbol{z_B}$—is denoted as:

$$\Delta \Pr(s = 0)_{\boldsymbol{z_A} \to \boldsymbol{z_B}} = \frac{\exp(\boldsymbol{z_A}\hat{\gamma})}{1 + \exp(\boldsymbol{z_A}\hat{\gamma})} - \frac{\exp(\boldsymbol{z_B}\hat{\gamma})}{1 + \exp(\boldsymbol{z_B}\hat{\gamma})} \tag{A.3}$$

Importantly, either of the above two quantities could be extended to provide for *posterior* probabilities of inflation (i.e. the probability of inflation conditional on an observation lying in the $Y = 0$ category). Such probabilities would more directly answer the question, "given that I observe a baseline outcome, what's the probability that it arises from either source?" (I thank an anonymous reviewer for this suggestion.) Comparable posterior probabilities have also been reported in Greene (2011) and Greene and Hensher (2010). An example of this *posterior* adjustment for equation (A.2) would be:

$$\begin{aligned} \Pr(s = 0|Y = 0) &= \frac{\Pr(\text{inflated}, Y = 0)}{\Pr(Y = 0)} \\ &= (1 - \Lambda(\boldsymbol{z}\hat{\gamma})) \div \left([1 - \Lambda(\boldsymbol{z}\hat{\gamma})] + \frac{\Lambda(\boldsymbol{z}\hat{\gamma})}{\sum_{j=1}^{J} \exp(\boldsymbol{x}\hat{\beta}_j)}\right) \end{aligned} \tag{A.4}$$

The second set of BIMNL predicted probabilities and probability changes corresponds to the probabilities of observing each multinomial choice outcome ($j = 0, 1, \ldots, J$) given specified values (or changes in) one's outcome stage covariates and one's inflation stage covariates. That is, these probabilities give an observation's aggregate probability of each choice outcome—after conditioning on that observation's predicted inflation probability $\Pr(s = 0)$. The first of these "global choice probabilities" captures the probability of observing one's baseline category outcome, and is presented immediately below:

$$\begin{aligned}\Pr(Y=0) &= [1-\Lambda(\boldsymbol{z}\hat{\gamma})] + \left(\frac{\Lambda(\boldsymbol{z}\hat{\gamma})\exp(\boldsymbol{x}\hat{\beta}_0)}{\sum_{j=1}^{J}\exp(\boldsymbol{x}\hat{\beta}_j)}\right) \\ &= [1-\Lambda(\boldsymbol{z}\hat{\gamma})] + \left(\frac{\Lambda(\boldsymbol{z}\hat{\gamma})}{\sum_{j=1}^{J}\exp(\boldsymbol{x}\hat{\beta}_j)}\right)\end{aligned} \tag{A.5}$$

where $\Lambda(\cdot)$ is the logistic c.d.f. For the non-baseline category responses (i.e. for $j = 1, \ldots, J$), the corresponding probability statement is:

$$\Pr(Y=j) = \left(\frac{\Lambda(\boldsymbol{z}\hat{\gamma})\exp(\boldsymbol{x}\hat{\beta}_j)}{\sum_{j=1}^{J}\exp(\boldsymbol{x}\hat{\beta}_j)}\right) \tag{A.6}$$

Next, I present the "global choice probability" changes in $\Pr(Y=0)$ and in $\Pr(Y=j)$ associated with a specified change in the values of the covariate vector $\boldsymbol{x}$ from $\boldsymbol{x_A}$ to $\boldsymbol{x_B}$. First for $\Delta\Pr(Y=0)$ under the assumption of $\boldsymbol{x}\equiv\boldsymbol{z}$, one can define this probability change as:

$$\begin{aligned}\Delta\Pr(Y=0)_{\boldsymbol{x_A}\rightarrow\boldsymbol{x_B}} &= \left[[1-\Lambda(\boldsymbol{z_B}\hat{\gamma})] + \left(\frac{\Lambda(\boldsymbol{z_B}\hat{\gamma})}{\sum_{j=1}^{J}\exp(\boldsymbol{x_B}\hat{\beta}_j)}\right)\right] \\ &\quad - \left[[1-\Lambda(\boldsymbol{z_A}\hat{\gamma})] + \left(\frac{\Lambda(\boldsymbol{z_A}\hat{\gamma})}{\sum_{j=1}^{J}\exp(\boldsymbol{x_A}\hat{\beta}_j)}\right)\right]\end{aligned} \tag{A.7}$$

while noting that, for the case where the modified covariate(s) in $\boldsymbol{x_A} \rightarrow \boldsymbol{x_B}$ are *not* found in $\boldsymbol{z}$, the above formula reduces to:

$$\Delta\Pr(Y=0)_{\boldsymbol{x_A}\rightarrow\boldsymbol{x_B}} = \frac{\Lambda(\boldsymbol{z}\hat{\gamma})}{\sum_{j=1}^{J}\exp(\boldsymbol{x_B}\hat{\beta}_j)} - \frac{\Lambda(\boldsymbol{z}\hat{\gamma})}{\sum_{j=1}^{J}\exp(\boldsymbol{x_A}\hat{\beta}_j)} \tag{A.8}$$

I can next present the comparable "global choice probability" change formula for $\Delta Pr(Y=j)$ when the modified covariate(s) in $\boldsymbol{x}\equiv\boldsymbol{z}$,

$$\Delta\Pr(Y=j)_{\boldsymbol{x_A}\rightarrow\boldsymbol{x_B}} = \frac{\Lambda(\boldsymbol{z_B}\hat{\gamma})\exp(\boldsymbol{x_B}\hat{\beta}_j)}{\sum_{j=1}^{J}\exp(\boldsymbol{x_B}\hat{\beta}_j)} - \frac{\Lambda(\boldsymbol{z_A}\hat{\gamma})\exp(\boldsymbol{x_A}\hat{\beta}_j)}{\sum_{j=1}^{J}\exp(\boldsymbol{x_A}\hat{\beta}_j)} \tag{A.9}$$

where the equivalent $\Delta Pr(Y=j)$ formula for the case where modified covariate(s) in $\boldsymbol{x_A} \rightarrow \boldsymbol{x_B}$ are *not* found in $\boldsymbol{z}$ is

$$\Delta\Pr(Y=j)_{\boldsymbol{x_A}\rightarrow\boldsymbol{x_B}} = \frac{\Lambda(\boldsymbol{z}\hat{\gamma})\exp(\boldsymbol{x_B}\hat{\beta}_j)}{\sum_{j=1}^{J}\exp(\boldsymbol{x_B}\hat{\beta}_j)} - \frac{\Lambda(\boldsymbol{z}\hat{\gamma})\exp(\boldsymbol{x_A}\hat{\beta}_j)}{\sum_{j=1}^{J}\exp(\boldsymbol{x_A}\hat{\beta}_j)} \tag{A.10}$$

The third and final set of BIMNL predicted probabilities and probability changes corresponds to the probabilities of observing each multinomial choice outcome $j = 0, 1, \ldots, J$ among a hypothetical non-inflated observation (i.e. for an observation whose $\Pr(s=1)$). I have defined these quantities of interest as "non-inflated choice probabilities" in the main paper. Immediately below, I first present the "non-inflated choice probability" formulas, which allow one to derive the probability of a non-inflated observation registering a particular choice category (over all others) given specified values on one's covariates. Doing so first for baseline category responses:

$$\begin{aligned}\Pr(\mathrm{Y}=0) \quad &= \frac{\exp\left(\boldsymbol{x}\hat{\beta}_0\right)}{\sum_{j=1}^{J}\exp(\boldsymbol{x}\hat{\beta}_j)} \\ &= \frac{1}{\sum_{j=1}^{J}\exp(\boldsymbol{x}\hat{\beta}_j)}\end{aligned} \tag{A.11}$$

and now for the non-baseline category responses (i.e. for $j = 1, \ldots, J$),

$$\Pr(\mathrm{Y}=\mathrm{j}) = \frac{\exp\left(\boldsymbol{x}\hat{\beta}_j\right)}{\sum_{j=1}^{J}\exp(\boldsymbol{x}\hat{\beta}_j)} \tag{A.12}$$

Next I present the "non-inflated choice probability" changes for $\Pr(Y = 0)$ and in $\Pr(Y = j)$, given a specified change in the values of the covariate vector $\boldsymbol{x}$ from $\boldsymbol{x_A} \rightarrow \boldsymbol{x_B}$. Note here that one does not need to set values for covariate vector $\boldsymbol{z}$. First presenting this quantity for the baseline outcome category, $\Delta\Pr(Y = 0)$,

$$\Delta\Pr(Y=0)_{\boldsymbol{x_A}\rightarrow\boldsymbol{x_B}} = \frac{1}{\sum_{j=1}^{J}\exp(\boldsymbol{x_B}\hat{\beta}_j)} - \frac{1}{\sum_{j=1}^{J}\exp(\boldsymbol{x_A}\hat{\beta}_j)} \tag{A.13}$$

and now for one's non-baseline choice categories, $\Delta\Pr(Y = j)$,

$$\Delta\Pr(Y=j)_{\boldsymbol{x_A}\rightarrow\boldsymbol{x_B}} = \frac{\exp(\boldsymbol{x_B}\hat{\beta}_j)}{\sum_{j=1}^{J}\exp(\boldsymbol{x_B}\hat{\beta}_j)} - \frac{\exp(\boldsymbol{x_A}\hat{\beta}_j)}{\sum_{j=1}^{J}\exp(\boldsymbol{x_A}\hat{\beta}_j)} \tag{A.14}$$