Procuring Peace after Prolonging War: International Institutions and the Durations of International Conflict and Post-conflict Peace*

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March 25, 2012

Abstract

How do international organizations (IOs) influence states' conflict behavior in the absence of centralized enforcement? This study develops and tests a theory about how IO membership helps solve the enforcement problems states face in the aftermath of a militarized conflict. It argues that joint membership in IOs that explicitly promote peaceful settlement of disputes improves enforcement conditions by increasing the costs of cease-fire violation in the long run. As a result, these IOs make a cease-fire more durable once the disputants agree to stop fighting. However, precisely because they expect longer peace after conflict, the member states have incentives to adopt tougher bargaining positions during conflict, causing a delay in reaching a cease-fire. A survival analysis that recognizes the interdependence between the durations of conflict and subsequent peace demonstrates that IO membership lengthens both the duration of conflict and the duration of subsequent peace.

*Paper prepared for presentation at the annual meeting of the American Political Science Association, September 1–4, 2011, Seattle, Washington. Earlier drafts of the paper were presented at the annual meetings of the International Studies Association, 2011, and the Midwest Political Science Association, 2011. I thank Songying Fang, Jesse Johnson, Ashley Leeds, and Bill Reed for their helpful suggestions and encouragements at various stages of the project. I am also grateful to comments from Phil Arena, Michael Butler, David Carter, Tobias Heinrich, Cliff Morgan, Meg Shannon, and Ric Stoll.

Introduction

There has been a growing body of research on the relationship between international organizations (IOs) and militarized interstate conflict. Many researchers have concluded that joint membership in IOs makes the onset of militarized conflict less likely. Nevertheless, there have been numerous violent conflicts fought between countries that belong to the same IOs. Despite its prevalence, less attention has been devoted to the potential effects of IOs on conflict that has already been initiated. Do IOs reduce the recurrence of conflict between member states? If so, how does this influence the disputants' behavior during conflict? This study develops and tests an argument about how IO membership influences the disputants' ability to reach and maintain a durable cease-fire agreement after conflict has begun.

Applying a model of bargaining and enforcement of international agreements to the case of cease-fire cooperation, I argue that IOs have two effects: they can make a cease-fire agreement more durable, but more difficult to achieve in the first place. More specifically, membership in IOs with a security mandate and resources to intervene can improve the enforcement conditions by making it more costly for the ex-belligerents to violate a cease-fire agreement. However, precisely because they expect longer peace after conflict, member states of these IOs will have incentive to hold out longer in conflict in hopes of obtaining better terms of a cease-fire, causing a delay in agreement.

To evaluate these predictions empirically, I develop a statistical model that recognizes that the duration of conflict and the durability of peace after conflict are a jointly determined outcome of international bargaining. The theoretical model suggests that the processes of bargaining and enforcement are interdependent, which means that the durations of militarized conflict and post-conflict peace are correlated. I use a copula function to characterize the joint distribution of the two correlated duration variables. Survival analysis of the data on conflict termination and recurrence from 1918 to 2004 provides support for the predictions.

The paper proceeds by laying out a theoretical argument about the effects of IO membership on the delay and durability of a cease-fire agreement in militarized conflict. The following section introduces the data and develops a research design to test the hypotheses. I then present the empirical findings. The final section concludes by discussing policy implications and the directions for future research.

IOs and Interstate Conflict

Scholars of international relations have devoted significant attention to the relationship between international institutions and militarized conflict. Earlier studies have reported that pairs of countries that share more memberships in IOs are less likely to become involved in militarized conflict (Oneal & Russett 1999, Oneal, Russett & Berbaum 2003, Russett & Oneal 2001, Russett, Oneal & Davis 1998). Subsequent studies have tried to provide specific causal mechanisms of conflict reducing effect of the IOs by drawing on bargaining theories (e.g. Boehmer, Gartzke & Nordstrom 2004). Bargaining models of war posit that militarized conflict occurs when bargaining fails to produce a more favorable outcome than war, and conflict is terminated when the disputant states reach some agreement that is better than continued fighting (Fearon 1995, Powell 2002, Reiter 2009). Building on the bargaining framework, recent empirical studies on IOs and conflict argue that joint membership in IOs reduces conflict by helping states overcome obstacles related to the bargaining process (Bearce & Omori 2005, Bearce, Floros & McKibben 2009, Boehmer, Gartzke & Nordstrom 2004, Haftel 2007, Hansen, Mitchell & Nemeth 2008, Mitchell & Hensel 2007, Pevehouse & Russett 2006, Shannon 2009, Shannon, Morey & Boehmke 2010).

While these studies contribute to our understanding of conflict processes, they have focused only on the onset, or the duration of conflict while ignoring what happens after the conflict is terminated. As a result, much remains to be learned about the effect of IO membership on the stability of peace after conflict. For example, does the conflict-reducing effect of IOs last after the fighting is terminated? If so, how does that affect states' behavior during conflict? If state leaders are rational and forward-looking, they will take into account what they think will happen in the future when they make decisions about their current course of actions. In other words, factors associated with the stability of post-conflict peace (such as IO membership) will also influence how the disputants fight a conflict before a cease-fire is reached between them. This calls into question the previous findings on IOs and conflict behavior that failed to take into account the disputants' expectation of future interaction.

In a similar vein, existing research on the duration of post-conflict peace has not fully explored how some exogenous conditions influence both the durability of a cease-fire *and* whether and when a cease-fire is achieved in the first place. Previous studies have focused on the effect of factors that are endogenously determined during conflict, such as institutional strength of a cease-fire agreement (Fortna 2003, 2004), the amount of informational asymmetries reduced in the battle fields (Werner & Yuen 2005), international mediation (Beardsley 2008, Gartner & Bercovitch 2006), types of political and military outcomes achieved by the disputants (Senese & Quackenbush 2003, Quackenbush & Venteicher 2008), or victor-imposed regime change at war's end (Lo, Hashimoto & Reiter 2008). While these factors are rightly of important considerations for scholars and practitioners of international politics, they are a consequence, not cause, of the expected difficulty of enforcing a cease-fire agreement in a given environment. As a result, much still remains to be learned about how the underlying enforcement conditions shape the likelihood and success of cease-fire cooperation. The challenge here is that simply looking for the correlates of successful enforcement of an agreement does not give us much leverage in pinpointing the determinant of cooperation, because the disputant states strategically choose whether and how to cooperate in the first place (Downs, Rocke & Barsoom 1996).

I address this problem by applying a formal model of international cooperation developed by Fearon (1998) that analyzes the bargaining process and the enforcement of an agreement jointly. The model consists of the initial stage where states bargain over the terms of cooperation, followed by the enforcement stage where states implement the the agreed terms in the face of temptations to renege. The key feature of the model is that the (expected) durability of a cooperative agreement in the second stage influences how long the states are willing to hold out in the first stage in hopes of obtaining better terms of cooperation. Although the model is not directly about militarized conflict, disputants in militarized conflict face similar incentives: the longer the disputants expect the cease-fire agreement to last, the longer they are willing to fight in hopes of obtaining better terms of a cease-fire. Then, if joint membership in IOs indeed makes post-conflict peace more stable, it may also have an unfortunate effect of prolonging conflict. Below, I discuss Fearon's model in detail and relate it to the specific subject of IO membership and cease-fire cooperation.

IOs, Bargaining, and Enforcement

The game captures strategic interaction between two states, 1 and 2, fighting over some flow of benefit worth one to each side per unit of time. As the states incur per-unit-time costs of conflict ($-c_i$, $i = 1, 2$) while fighting, they have incentive to agree on a cease-fire and split the disputed good according to some division. But, the disputants disagree over how to divide the disputed good. In addition, the anarchic nature of the international system allows either state to deviate from the agreed division, even if it is reached in bargaining. Therefore, the disputants face two distinct problems as they go through two consecutive stages. In the initial bargaining stage, they must agree on which of the potential cease-fire agreements to implement after they stop fighting. In the subsequent enforcement stage, they must ensure that the agreed deal is implemented in the face of temptations to renege. In the following two sections, I discuss each of the two stages and derive testable hypotheses based on the model's comparative statics. The two stages are analyzed in the backward order because the disputants' bargaining behaviors in the initial stage are determined by their anticipation of what will happen after they agree to stop fighting.

IOs and Enforcement of a Cease-fire

The enforcement stage begins when the disputants reach some agreement about the division of the disputed good. Let $x \in [0, 1]$ be such an agreement that gives x to state 1 and $1 - x$ to state 2 per unit of time. The two states now face the problem of ensuring compliance with x by

		State 2	
		C	D
State 1	C	$x, 1-x$	$-b, a$
	D	$a, -b$	$-c_1, -c_2$

Figure 1: Per-unit-time payoffs of enforcement stage

both sides. If both sides deviate from the agreed cease-fire, they go back to the bargaining stage, where neither gets the disputed good and both pay the costs of conflict, c_i. If one state deviates from the agreement while the other state complies with it, the defector can get a payoff greater than the agreement payoff ($a > x$ and $a > 1 - x$). The compliant state is assigned the "sucker" payoff ($-b < -c_i$) during this period, giving it incentive to defect as well. Therefore, a unilateral defection is profitable for the defector as long as the other state does not respond with retaliatory defection in future rounds for some period of time. Such a delay in response is captured by a response delay parameter, $\Delta > 0$, such that if one state unilaterally defects at time t, the other state is unable to detect or to respond to this defection until $t + \Delta$. Figure 1 describes the per-unit-time payoffs of the enforcement stage, where states choose between C (comply) or D (defect). As this game is played continuously over time, the disputants evaluate streams of future per-unit-time payoffs according to a constant discount rate $r > 0$, meaning that receiving x over d consecutive units of time into the future is worth $e^{-rd}x$ now. Smaller discount rates indicate that states discount future payoffs less, making the shadow of the future longer.

A cease-fire agreement is upheld in the enforcement stage as long as both sides can make a credible threat of retaliation. It is the disputants' ability and willingness to punish a unilateral defection that guarantees mutual compliance with the cease-fire. To see this more precisely, Fearon (1998) shows that a simple and severe grim-trigger enforcement scheme can maintain cooperation when the following condition holds:

$$r\Delta \leq \min\left\{ \ln \frac{a + c_1}{a - x}, \ln \frac{a + c_2}{a - (1 - x)} \right\}. \tag{1}$$

This condition suggests that an agreement is easier to enforce the longer the shadow of the future (smaller r); the quicker the detection of a unilateral defection (smaller Δ); the lower the per-unit-time benefit of unilateral defection (smaller a); the greater the costs of non-agreement (greater c_i).

I argue that membership in IOs improves the enforcement condition in the context of cease-fire cooperation by extending the shadow of the future. The shadow of the future is interpreted as players' assessment of the probability of continued interaction under the same payoff structure, or how much values players assign to future payoffs.[1] Longer shadows of the future make it more costly for the states to defect unilaterally from the agreed cease-fire, because the opponent can more credibly threaten to punish the defection in future rounds of interactions.[2] The improved prospect for enforcement, in turn, further extends the shadow of the future by encouraging the states to weigh future payoffs more relative to present payoffs.

How does joint participation in IOs extend the shadow of the future of member states? First, an important aspect of membership in IOs is that it fosters an expectation that the member states will interact with each other for a very long time. Unless states have some intention to continue interacting with other members, they will not form or join an IO in the first place. In addition, some IOs are equipped with institutional devices to make unilateral violation of a cease-fire less attractive even in the short run, which will in turn stabilize the relationship in the long run. Although many IOs have been created in hopes of promoting peace, not all institutions have a security mandate and resources to help states enforce the agreed cease-fire. I thus focus on *pacific settlement IOs*, defined as international institutions (including both intergovernmental organizations and multilateral treaties) that call for peaceful settlement of disputes among members (Hensel 2005). [3] It is these IOs that can monitor compliance behavior with respect to the agreed cease-fire, facilitate

1 Having longer shadow of the future is equivalent to having smaller r, or lower discount rate. Bearce, Floros & McKibben (2009) and Mitchell & Hensel (2007) make a similar assumption in their analyses of IO membership and international cooperation.

2 Of particular importance here is the stability of payoff ordering between b and c_i. This is because the credibility of enforcement threats depends on the willingness of the other state to bear the costs of mutual defection ($-b < -c_i$) in case a unilateral defection is detected.

3 The analysis below thus excludes institutions that are primarily about economic cooperation, or other non-security issues. There have been 51 pacific settlement IOs in the world. Examples of these institutions include global IOs, such as the United Nations, and Non-aligned Movement, as well as regional institutions, such as the African Union, League of Arab States, and the Organization of the American States.

communication among the ex-belligerents, and increase opportunity costs of conflict.

There are numerous historical case examples where pacific settlement IOs take proactive actions to help disputants enforce a cease-fire. For example, in the 1991 crisis in Yugoslavia, a handful of international organizations, including the United Nations (UN), the European Community (EC), Conference for Security & Cooperation in Europe, and North Atlantic Treaty Organization (NATO), were involved in the cease-fire process. After several violations of the cease-fire agreements brokered by these IOs, the UN, EC, and NATO imposed economic sanctions and launched air strikes on Yugoslavia to enforce the peace agreement (Hufbauer, Schott, Elliott & Oegg 2007, Woodward 1995). Regional IOs outside the Europe have also been actively involved in the conflict management processes in disputes between their member states. For example, the League of Arab States (LAS) was highly involved in the war between North and South Yemen in 1979. The mediation efforts sponsored by the LAS culminated in a cease-fire between the disputants on March 3, 1979. As fighting continued despite the cease-fire agreement, on the 5th, the LAS further adopted a resolution to supervise the implementation of the cease-fire and the process of normalizing relations between the disputants. Then, the withdrawal of the troops by both sides was completed in several days under the auspices of the LAS (Bidwell 1983).

In addition to these instances where actions of IOs were directly observable, there are many more instances where pacific settlement IOs exercise an implicit pressure on the member states. Recent formal works show that IOs may be able to influence states' conflict behavior even when they are not directly involved in the dispute processes (e.g. Fang 2010). Moreover, because of the explicit treaty obligation to resolve their disputes peacefully, member states of these IOs find it more costly to make a credible threat to revise the status quo distribution of benefit with force, which adds an additional stability to the agreed terms of a cease-fire. I expect that the greater the number of pacific settlement IOs that the disputants both belong to, the greater these effects will be. This logic suggests that membership in pacific settlement IOs has an effect of stabilizing the peace after a conflict is terminated between member states.

Hypothesis 1 (IO effect in the enforcement stage) *The disputants will have a longer duration of peace after a termination of militarized conflict when they share memberships in more IOs that explicitly promote pacific settlement of conflicts among member states.*

IOs and Bargaining for a Cease-fire

Although pacific settlement IOs make a cease-fire agreement more durable in the long run, improved enforcement can also make a cease-fire agreement more difficult to achieve. In other words, when states anticipate that they will be stuck with the agreed terms for a long time, they have incentive to fight longer to obtain better terms of a cease-fire before agreeing to it. This can result in a costly delay in a cease-fire. To explore the effect of longer shadows of the future on the disputants' incentive to terminate conflict, I now turn to the initial bargaining stage that precedes the enforcement stage.

Prior to the enforcement stage, the states are in dispute over two possible deals (divisions of the disputed good) that both parties prefer to continued fighting. The problem is that the disputants disagree over which one to implement in the subsequent enforcement stage. Denote State i's preferred division by $x_i \in [0, 1]$, where $x_1 > x_2$. I assume that the two feasible divisions are fixed for reasons exogenous to the model, such that it is prohibitively costly for the disputants to come up with a new offer. The disputants resolve this disagreement by playing a continuous time war-of-attrition game where the player who quits fighting first has to concede the preferred division to the other.

A pure strategy for state i is the quit time $t_i \geq 0$ that specifies how long it will hold out in hope of getting the better agreement. When either state quits fighting, a cease-fire agreement is struck, and the states move to the enforcement game where the winner's preferred division is implemented. For example, if $t_1 > t_2$, the enforcement game begins at time t_2 with per-unit-of-time payoffs of mutual compliance is $(x_1, 1 - x_1)$. While fighting, states incur per-unit-time costs of non-agreement, $c_i > 0$, the magnitude of which determines the relative power of states on the disputed issue. That is, states with smaller c_i will be willing to fight longer for a better deal, on average. Ex ante, the states do not know who is stronger (have lower costs of non-agreement) and thus willing to hold out longer in bargaining. [4] However, after they enter into a costly war of attrition, the states can credibly signal their "strength" by bearing the costs of non-agreement. Fearon (1998) shows that in a symmetric Bayesian equilibrium in the bargaining phase, the ex ante expected time until agreement is [5]

$$\bar{t} = \frac{\ln 8 - 1}{r} \simeq \frac{1}{r}. \tag{2}$$

This means that the longer the shadow of the future (smaller r), the longer the duration of costly conflict until agreement will be.

While condition (1) suggests that the longer shadow of the future facilitated by IO membership contributes to better prospect of enforcement, condition (2) implies that it is a double-edged sword. If disputants expect more durable peace after agreeing to terminate conflict, they have incentive to adopt tougher bargaining strategies and hold out longer in conflict (the quit time t becomes greater) so as to extract better terms of peace. On the other hand, if the disputants expect that post-conflict peace will be fragile because of greater risk of changes in the payoff structure, they will quickly terminate conflict with little intention to follow through on the cease-fire agreement for very long. [6] I thus have the following testable hypothesis.

Hypothesis 2 (IO effect in the bargaining stage) *The disputants will have a longer duration of conflict when they share memberships in more IOs that explicitly promote pacific settlement of conflicts among member states.*

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[4] Fearon (1998) specifies that the states know their own costs, but that they know only the distribution of their opponent's costs. For convenience, it is assumed that the costs are randomly drawn from uniform distributions on the interval $[1, 2]$.

[5] Fearon (1998) also solves the complete information version of the game. Although the structure of the game is common knowledge, costly fighting occurs in equilibrium as states employ mixed strategies.

[6] Note that a mere expectation of future changes in the payoff structure can undermine mutual compliance because of the "last-period effect" in repeated games.

Research Design

To test the hypotheses, I analyze historical data on the duration of violent conflict and the subsequent duration of peace after the termination of active fighting. These two outcome variables are constructed using information from the International Crisis Behavior (ICB) project. I use Hewitt's (2003) dyadic version of the data set.[7] The initial analysis includes those crises that involve actual fighting, while excluding non-violent crises.[8] There are 435 crisis-dyads from 1918 to 2004, involving 234 unique international crises. [9] All the crisis observations in the sample have been terminated by the end of 2004.

I operationalize the first-stage outcome variable, duration of conflict until a cease-fire agreement, as the time elapsed between the date of initiation and date of termination of international crises in a dyad. A "cease-fire" here does not necessarily mean a complete resolution of the underlying contentious issues between the disputants nor a formal armistice treaty signed by the disputants; it is simply defined as a cessation of active fighting. Therefore, a cease-fire here can be informal (e.g., oral declaration or tacit understanding). As long as neither side of the former belligerents is engaged in active fighting, the disputants are considered to be in the post-conflict peace phase, be it fragile and short-lived. This operationalization reflects the notion of cooperation in the theoretical discussion above, where cooperation represents an absence of unilateral and mutual defection in the face of short-term benefit of reneging.

With this coding rule, there are 427 crisis-dyads where the disputants stop active fighting and proceed to the post-conflict peace phase. In the remaining 8 ($= 435 - 427$) cases, the defeated country ceases to exist as an independent state because of annexation by the victor. [10] I treat these 8 observations as right-censored. That is, I allow the "true" duration of conflict until a cease-fire

7 I expanded the data set by adding newer ICB cases, based on the latest version of the actor-level and crisis-level data (version 10, July 2010).

8 That is, I include those ICB crises that reach at least the second level of violence ("Minor clashes") in the ICB data set. I relax this assumption later in the paper.

9 There are several overlapping crises in the data set, where a next crisis begins before the previous one is terminated. In these cases, I combine the two crises into a single crisis observation.

10 All of these incidences are from the crises that happened during the Second World War.

(i.e., the duration that would have been observed if the defeated state had not lost independence) for these cases to be at least as long as the observed (censored) duration. [11] For the uncensored 427 observations, the duration of conflict ranges from 1 to $1,462$ days in the sample, with median 110 and mean 198 days.

The second-stage outcome variable, the durability of enforcement phase following a cease-fire agreement, is operationalized as the time elapsed between the termination and recurrence of international crisis in a dyad. Among those 427 post-crisis-dyad observations, a crisis has recurred in 168 cases (39.3%), for which the uncensored duration of post-conflict peace is fully observed. Since post-conflict peace has not broken down in the remaining 259 post-crisis-dyads as of 2004, the peace duration for these cases is treated as right-censored at the end of 2004. Uncensored duration of post-conflict peace is highly skewed to the right, ranging from 8 to $17,916$ days (49 years), with a median of $1,473.5$ days (4 years) and a mean of $2,825$ days (8 years). To incorporate time-varying covariates, those crisis-dyads and post-crisis-dyads that span multiple years are duplicated accordingly, yielding a total of 675 crisis-dyad-year observations for conflict duration and $7,706$ observations for peace duration.

Explanatory Variables

Data on the key explanatory variable, membership in pacific settlement IOs, are taken from the Multilateral Treaties of Pacific Settlement (MTOPS) data set of the Issue Correlates of War (ICOW) project (Hensel 2005). The MTOPS data set contains annual observations of state membership in qualifying IOs from 1816 to 2004. To be qualified as such, pacific settlement IOs must call explicitly for peaceful settlement of disputes among signatories, rather than simply mentioning the desirability of peace. The MTOPS data set is limited to those IOs that have more than five member states. To capture the effects of IO membership on the disputants' shadow of the future, I count the number of shared membership in these IOs in a dyad. The variable ranges from 0 to 11, with a median of 2.

[11] Alternatively, I also run all the analyses below by dropping these 8 observations entirely from the data set. The results are qualitatively the same.

I also include an interaction term between the IO membership and the analysis time, which allows for the effects of IO membership to change over time. This is to account for the disputants' diminishing uncertainty about each other's resolve to bear the costs of conflict. In the initial phase of conflict, there is much to learn about the other side's private information about their resolve to bear the costs of conflict. In fact, this uncertainty about each other's costs of conflict is the very reason they enter the conflict in the first place. Fighting allows the disputants to learn about each other's private information, and to update their beliefs accordingly (Slantchev 2003). As time goes by, however, the amount of additional information to be learned will necessarily shrink. This is where the expectation of the future begins to loom large. Therefore, IO membership will have greater effect in the later phases of conflict when the information asymmetries are relatively leveled down.

A set of control variables that may confound the relationship between IO membership and conflict are also included in the analysis. *Contiguity* is a binary indicator measuring whether or not the disputants share borders; *Major Power* takes the value of 1 if a crisis-dyad involves at least one major power; *Capability Balance* is the ratio of the stronger state's military capability over the sum of capabilities in a dyad; *Joint Democracy* is a binary variable coded as 1 if both states in a dyad have a 6 or higher polity score (democracy minus autocracy score), and 0 otherwise; and the *Year of Initiation* measures the year in which the crisis breaks out. [12]

Statistical Model

In order to conduct a systematic analysis of the data, we need a statistical model that can appropriately capture the theoretical insight of Fearon's (1998) formal model that the duration of conflict and the duration of peace are a jointly determined outcome of bargaining. The model shows that enforcement conditions affect not only how long the agreed cease-fire will last once it is reached but also how long the disputants are willing to hold out before a cease-fire. Then, if some unmeasurable enforcement conditions influence both processes, the "residual" duration of conflict

[12] Data on *Contiguity*, *Major Power*, and *Capability Balance* are taken from the Correlates of War data set. Data on democracy score is from the Polity IV data set.

(conflict duration conditional upon all the observables) will be correlated with the residual duration of peace. For example, suppose some dyads have shorter shadows of the future than other dyads, because of some unmeasurable variables. Conditions (1) and (2) imply that these dyads would experience shorter delays in reaching an agreement, but the post-conflict peace in these dyads would necessarily be fragile. This yields a positive correlation between the two residual durations. Similarly, a failure to measure and control for costs of conflict generates a negative correlation in the error term. [13] To the extent that common unmeasured factors systematically influence both bargaining and enforcement stages, analyzing one stage while ignoring another will lead to bias in our inferences about the determinants of both processes. [14]

To cope with this, I construct a unified statistical model of bargaining and enforcement that controls for unobservable enforcement conditions influencing the delay and durability of a cease-fire agreement. The proposed solution is to jointly estimate the durations of conflict and post-conflict peace while accounting for the correlation between the two. The model is a generalization of the parametric univariate duration model in the sense that it can readily resolve the issues of right-censoring and duration dependence as in other duration models.

Let $i = 1, \ldots, n$ denote a pair of states in a crisis, the observation unit I call crisis-dyad. The goal is to make inferences about the duration of the initial bargaining stage, or time until crisis-dyad i agree to stop fighting, T_{ci}, and the duration of peace enforcement stage, or time until crisis-dyad i resumes conflict after a cease-fire, T_{pi}. For crisis-dyads where a cease-fire agreement is reached, we observe the actual value of t_{ci} as a realization of the underlying random variable T_{ci}. For right-censored cases, we only observe the right-censoring point, t^0_{ci}. I define a binary cease-fire indicator, A_i, that takes on the value of 1 in crisis-dyads with a cease-fire, and 0 in the right-censored crisis-dyads. Then, for crisis-dyads where a cease-fire is reached and post-agreement peace has failed by the end of observation period, we observe the actual value of peace duration t_{pi} as a realization of

[13] Some dyads with unobserved greater costs of non-agreement should experience *shorter* durations of conflict and *longer* durations of post-conflict peace, on average, than other dyads with similar observable characteristics.

[14] Furthermore, there may be some factors *outside the theoretical model* that cause a correlation between the two durations. For example, the war-weariness argument suggests that there will be longer peace after longer conflicts because the disputants become weary of wars after a long fighting.

the underlying random variable T_{pi}. For crisis-dyads where crisis was terminated and post-conflict peace has not failed yet, we only observe the right-censoring point, t^0_{pi}. I define another binary indicator, B_i, that takes on the value of 1 in post-agreement crisis-dyads with recurrence, and 0 otherwise.

With these notations, the likelihood function is given as follows:

$$\mathcal{L} = \prod_{i=1}^{n} \Pr(T_{ci} > t^0_{ci})^{(1-A_i)} \Pr(T_{ci} = t_{ci} \cap T_{pi} > t^0_{pi})^{A_i(1-B_i)} \Pr(T_{ci} = t_{ci} \cap T_{pi} = t_{pi})^{A_i B_i}. \quad (3)$$

The first component of the likelihood function, $\Pr(T_{ci} > t^0_{ci})$, represents the likelihood contribution from observations that are right-censored during the conflict phase, the second component, $\Pr(T_{ci} = t_{ci} \cap T_{pi} > t^0_{pi})$, corresponds to those observations that are right-censored during the peace phase after the disputants reached an agreement to stop fighting. The last component, $\Pr(T_{ci} = t_{ci} \cap T_{pi} = t_{pi})$, is for those observations where disputants reach an agreement once and then fighting resumes.

To specify the likelihood function (3), I begin by characterizing the univariate marginal distribution of the two random variables, T_{ci} and T_{pi}. I allow the durations of conflict and peace to be conditioned on vectors of covariates, z_{ci} and z_{pi}, respectively. Then the hazard rates governing the two durations are specified as $\lambda_{ci} = \exp(-z_{ci}\beta_c)$ for conflict duration and $\lambda_{pi} = \exp(-z_{pi}\beta_p)$ for peace duration, where β_c and β_p are vectors of the coefficient parameters. Using the flexible Weibull specification, the univariate density function $f(t)$, the survivor function $S(t)$, and the distribution function $F(t)$ are each given as a function of λ and the shape parameter σ. [15] The shape parameter determines whether the risk of "failure" event (i.e., conflict termination in the bargaining stage or conflict recurrence in the enforcement stage) is increasing ($\sigma > 1$), decreasing ($\sigma < 1$), or constant ($\sigma = 1$) over analysis time.

The challenge here is to characterize the joint distribution of two duration variables, which is

[15] I chose the Weibull model because of its flexibility and simplicity. One limitation of the Weibull specification, however, is that it does not allow for non-monotonic change in hazard. As a robustness check, I also estimate a log-logistic model that allows for non-monotonicity (but does not allow for a monotonic increase) in hazard. Statistical results are qualitatively the same, and Vuong's (1989) test of non-nested models does not reject the null hypothesis that the two models are equally good.

necessary to specify the second and the third components of the likelihood function (3). I utilize a copula function to derive a joint distribution from the two univariate marginal distributions. A copula is a function that binds together two or more univariate marginal distributions of known form to produce a new joint distribution (Trivedi & Zimmer 2005). As we have two marginal distributions, the association between the two marginals is represented by a single association parameter, θ, which captures the correlation between the residual durations of conflict and subsequent peace. The appendix provides the derivation of the statistical model in greater details.

Estimation Results

Using the data described so far, I maximize the likelihood function (3) with respect to the coefficient (β_c, β_p), the shape (σ_c, σ_p), and the correlation (θ) parameters jointly in full information maximum likelihood. Table 1 reports the maximum likelihood estimates of each parameter. Since there are two dependent variables, there are two sets of coefficients: in the top panel, the first numerical column shows the estimates of the coefficients for the conflict duration (β_c), and the second numerical column shows those for the peace duration (β_p). These coefficients are in the accelerated failure time metric, which means that positive estimates are associated with longer durations and negative estimates are associated with shorter durations.

The second panel in the table shows the estimates of the auxiliary parameters. The Weibull shape parameters for conflict (σ_c) and peace (σ_p) durations are greater than (and statistically distinguishable from) 1. This indicates that the baseline hazard of conflict termination and that of recurrence are both increasing over time. The correlation parameter (θ) measures the interdependence between the two residual durations, and assumes values between -1 (perfect negative correlation) and $+1$ (perfect positive correlation). The estimated θ is negative, and statistically distinguishable from zero. This result suggests that the residual durations of conflict and post-conflict peace are negatively correlated. As discussed above, one possible explanation for a negative correlation is our inability to measure the costs of conflict.

[Table 1 About Here]

As for the effects of IO membership variables on conflict and peace durations, we obtain negative estimates for the raw IO variables and positive estimates for the interaction terms between IO membership and the analysis time in both stages. Since this is a non-linear model and there are interaction terms, it is not immediately clear what these estimates mean substantively. I thus calculate and plot probabilities of continued fighting and those of conflict recurrence for different values of the IO variable and observation time, while setting all the other variables at their median values. Figure 2 illustrates the estimated effect of IO membership on the duration of conflict. The horizontal axis shows time (measured in months) and the vertical axis represents the survival probability of conflict, or the probability of continued fighting beyond time t. Greater probabilities of conflict survival mean that the duration of conflict is longer. The vertical ticks at the bottom of the graph show the observed distribution of conflict in the data set. For this figure, I compare two values of the IO membership variables. The first is 1, which corresponds to the case where the disputants are members of one pacific settlement IO. The other is 3, which is the case where disputants are members of three IOs. These values correspond to the 25th and 75th percentile values of the IO membership variable in the data set. The curve in light gray shows the estimates when membership is set equal to 1, with 95 % confidence intervals, and the dark gray curve shows them when the membership is 3. These curves show the estimated survivor function $\Pr(T_c > t_c)$, conditional on $1,000$ simulated values of the parameters and substantively interesting covariate profiles. [16]

We can see form Figure 2 that the overall effect of IO membership is to increase the duration of conflict, lending support for Hypothesis 1. Membership in pacific settlement IOs indeed makes a violent conflict more difficult to terminate. At the same time, we can also see that, in the initial days of conflict (when $t < 1.5$ month), the IO membership has a small negative impact on the probability of continued fighting, but as time passes the effect becomes positive and grows larger. This observed pattern is consistent with the following interpretation. First, members of pacific

[16] Simulated parameters are generated by randomly drawing $1,000$ values from a multivariate Normal distribution characterized by the maximum likelihood estimates of the parameters and the estimated variance-covariance matrix (King, Tomz & Wittenberg 2000).

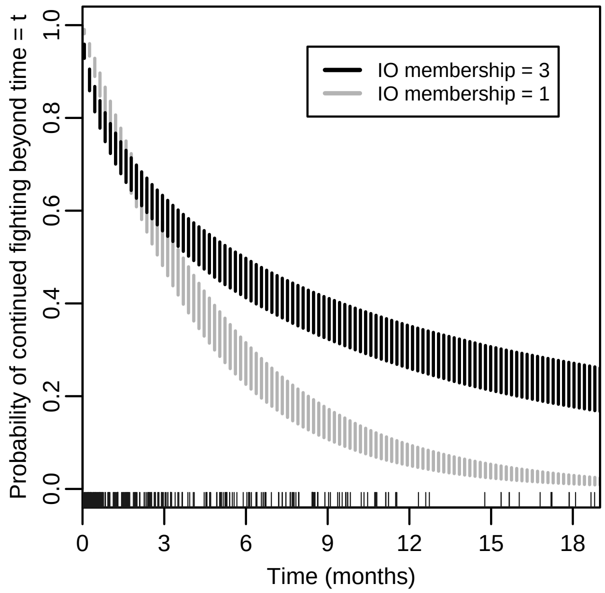

Figure 2: Substantive effects of IO membership on conflict duration

(*Notes.* This figure shows the estimated impacts of IO membership on conflict duration along with 95 % confidence bands. The vertical axis shows the predicted probabilities of continued fighting beyond time $= t$, and the horizontal axes show the analysis time in months).

settlement IOs have a better chance of terminating conflict in the very initial phases of conflict where the disputants face greater uncertainties about the likely outcome of conflict. This is perhaps because the informational gap between the disputants is smaller among IO members than among non-members. However, once a conflict "survives" the first two months or so, the informational asymmetry between the disputants has already diminished through fighting. This is when the disputants' concerns about enforcement conditions in the future begin to dominate. In other words, beyond this point, the improved enforcement conditions in the future as a result of membership in pacific settlement IOs start to lead to longer conflict.

Figure 3 shows the estimated substantive effects of IO membership on the durability of a ceasefire after a conflict. The vertical axis represents the estimated conditional probabilities of conflict recurrence beyond time $= t$, given the median duration of conflict. Smaller probabilities of conflict recurrence mean that the duration of post-conflict peace is longer. Again, the curve in light gray corresponds to the case where IO membership is 1, and the dark gray curve to the case where the

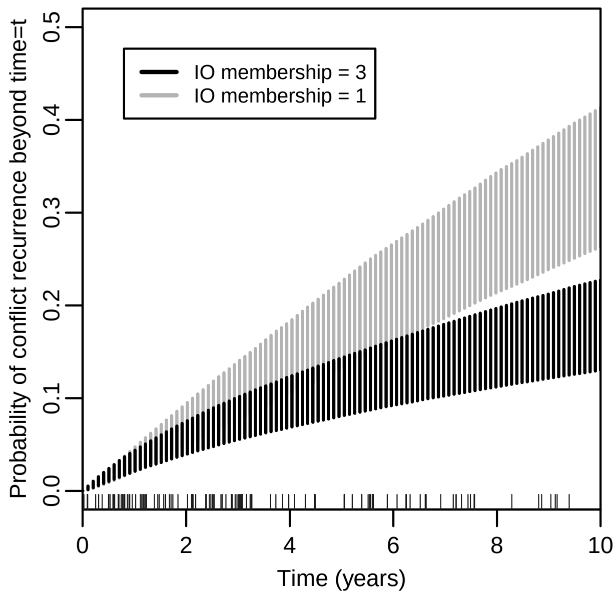

Figure 3: Substantive effects of IO membership on peace duration

(*Notes.* This figure shows the estimated impacts of IO membership on post-conflict peace duration, along with 95 % confidence bands. The vertical axes show the predicted conditional probabilities of conflict recurrence beyond time $= t$, and the horizontal axes show the analysis time in years. The vertical ticks at the bottom of the graph show the observed (i.e., uncensored) distribution of peace duration in the data set).

membership is 3. The figure shows that an increase in IO membership decreases the probability of recurrence, lending support for Hypothesis 2. We can also see that the pacifying effect of IO membership is initially small in the immediate aftermath of a conflict, but it becomes greater over time. Specifically, the conflict reducing effect is statistically indistinguishable from zero in the first three or four years after the termination of conflict, but it achieves statistical significance after that. [17]

Robustness Checks

The statistical findings provide evidence that the theorized causal mechanisms are indeed at work in historical data. As a robustness check, I examine the sensitivity of the initial results to several assumptions I make in analyzing the data. Table 2 reports results from four additional tests along

[17] Although it appears that the confidence intervals of the two curves overlap in the first six years or so, the *difference* between the two curves is statistically distinguishable from zero after about 3.5 years.

with the baseline estimates discussed above.

[Table 2 About Here]

First, we will see the consequences of ignoring the correlation between the durations of conflict and peace. Column (1) in Table 2 shows the results from a naive model that assumes independence between the conflict and peace durations. [18] Compared with the interdependent model, the naive model fits the data poorly, as the maximized log likelihood value from this model (-4289) is significantly smaller than that from the interdependent model (-4188). [19] In addition, there are several major differences in direction, size, and the statistical significance of the estimates. For example, the estimated Weibull shape parameter for the peace duration is smaller than 1 (indicating negative duration dependence) in the naive model, whereas it is greater than 1 (indicating positive duration dependence) in the interdependent model. This means that the naive model fails to capture the shape of underlying probability of conflict recurrence trending over time.

Second, I consider the possibility that the observed conflict-prolonging effect and/or peace-stabilizing effect of IOs depend on the types of international conflict included in the sample. Some may argue that IO membership has little effect on serious conflict that involves large scale violence while it may affect small scale disputes involving verbal threats or non-violent displays of force. Column (2) in Table 2 shows the results from an analysis that focuses on conflicts that involve serious military clashes. This model excludes those conflicts that do not reach the third level of violence ("Serious clashes") in the ICB data set. [20] The statistical results show that the estimated values of the important parameters are qualitatively the same as those obtained from the baseline model, indicating that the same dynamics emerge even when we focus exclusively on high-intensity conflicts. I also consider the possibility that excluding non-violent conflict biases the initial results. An analysis shown under column (3) includes those conflicts that do not reach the

[18] This is equivalent to estimating two duration models separately, one for conflict duration and another for peace duration.

[19] Akaike's information criteria for the naive model and the interdependent model are 8613 and 8414, respectively, showing that the interdependent model is better.

[20] The analysis thus includes the duration of conflict for 297 crisis-dyad observations and the duration of subsequent peace for 289 post-crisis-dyad observations, which yields a total of $5,907$ time-varying observations.

second level of violence ("Minor clashes"), along with more violent conflicts. Again, the results are qualitatively the same as the original results. Model (4) looks at those conflicts that occurred after the second World War to see if the findings are driven by the existence of the United Nations. Once again, the estimation results underline the same pattern.

Finally, I examine the threats to inference caused by potential sample selection effect. The set of international conflicts analyzed in this study may not be an appropriate sample to explore the effects of IO membership on the durations of conflict and post-conflict peace if there is a systematic relationship between IO membership and conflict occurrence. To address the issue of nonrandom sample selection, I analyze a three-stage model that incorporates the initial selection process into the interdependent duration model. For this analysis, we must first identify a set of observations that are at risk of experiencing (but do not necessarily experience) militarized conflict. I consider all pairs of countries that have competing territorial claims with one another, as identified by Huth & Allee (2002). Territorial disputes are particularly prone to escalate to violent conflict, but not all dyads with competing territorial claims experience conflict. The data set spans the time frame from 1919 to 1995. In this time period, there are 249 dispute-dyad observations, where an ICB crisis breaks out in 131 of the cases.

The model's first stage predicts the selection into conflict, which is specified as duration of peace after the initiation date of territorial claims and before the outbreak of violent conflict. [21] Once there is a conflict, the second stage predicts the duration of conflict given selection, and the third stage predicts the duration of post-conflict peace given selection and the duration of conflict. [22] The estimation is conducted by combining two interdependent duration models. That is, I first estimate the initial two stages jointly, and then estimate the latter two stages jointly by treating the second stage estimates as given. Table 3 displays the estimation results.

21 Those dispute-dyads that do not experience violent conflict are treated as right-censored. There are 118 ($= 249 - 131$) such observations in the data.

22 Although the other aspects of the research design are similar to the previous models discussed above, there are several differences in the measurement of covariates. First, the *Year of Initiation* variable in this model measures the year in which the territorial claim begins. Second, the cut point for the binary *Joint Democracy* in this model is 0, instead of 6. Since there are not many democratic dyads that experience territorial claims, the variable has little variation if we use 6.

[Table 3 About Here]

The estimated coefficients from the second and third stages show that IO membership has the hypothesized conflict-prolonging effect and peace-stabilizing effect even after accounting for the selection into conflict. We can thus conclude that the initial results are not an artifact of selection bias. Moreover, the estimated coefficients from the first stage show that IO membership has no discernible effects on the duration of peace before conflict. In other words, IO membership does not seem to be a good predictor of whether and when potential disputants experience violent conflict.

Conclusions

States create a variety of international institutions to help them achieve cooperation with other states in the absence of centralized enforcement. These institutions, once in force, alter the strategic environment in which states interact with others. To explore how IOs shape the member states' conflict behavior, this study applies a model of bargaining and enforcement to the case of cease-fire cooperation. The theoretical discussion suggests that the conflict-reducing effect of IOs manifests itself after the conflict is terminated. It maintains that membership in IOs that have an explicit obligation for member states to resolve disputes will make a conflict-ending settlement more durable. But, the improved enforcement, in turn, can encourage the disputants to stick to tougher bargaining positions, prolonging the duration of costly fighting preceding a cease-fire. I evaluate these arguments empirically, and find support for the hypotheses.

This research not only contributes to the scholarly debate about institutional influence on states' conflict behavior but also provides important implications for policy makers. Given that the great proportion of international conflicts have been fought between the same pairs of countries many times, the finding that IO membership reduces conflict recurrence suggests that extending the membership in pacific settlement IOs to these countries may be one way to address this problem. At the same time, we should also be aware of the danger in doing so. That is, if states anticipate that they will be stuck with the agreed terms of peace in the long run once they stop fighting, they have less incentive to terminate conflict. Once we recognize this trade-off, we can begin to think about possible ways to mitigate it. This recognition becomes of critical importance when we are to evaluate the effectiveness of various policy tools, such as third-party mediation, that have received significant attention in previous studies of militarized conflict. More specifically, research on third-party conflict management will be benefitted by taking into account the dual effects of IO membership demonstrated in this study in order to devise an optimal way to intervene in conflict.

Appendix: Derivation of the Interdependent Duration Model

I begin by characterizing the univariate marginal distribution of the random variables, T_{ci} and T_{pi}. Using the Weibull specification, the univariate density function $f(t)$, the survivor function $S(t)$, and the distribution function $F(t)$ are given as:

$$
\begin{aligned}
f(t) &\equiv \Pr(T = t) = \lambda\sigma\,(\lambda t)^{(\sigma-1)} \exp\left(-\lambda t\right) \\
S(t) &\equiv \Pr(T > t) = \exp\left(-\left(\lambda t\right)^{\sigma}\right) \\
F(t) &\equiv \Pr(T < t) = 1 - S(t)
\end{aligned}
$$

where λ and σ are as defined in the text. Then, I characterize three parts of the likelihood function in turn.

$$
\mathcal{L} = \prod_{i=1}^{n} \Pr(T_{ci} > t_{ci}^{0})^{(1-A_i)} \Pr(T_{ci} = t_{ci} \cap T_{pi} > t_{pi}^{0})^{A_i(1-B_i)} \Pr(T_{ci} = t_{ci} \cap T_{pi} = t_{pi})^{A_i B_i}. \quad (3)
$$

Specification of the first component of the likelihood function (3) is trivial, since this is just a univariate survivor function evaluated at t_{ci}^{0}. The next step is to calculate the second and the third component of the likelihood function, both of which are bivariate joint distributions of T_{ci} and T_{pi}. This is done by utilizing a copula function. A copula is a function that parameterizes the dependence between univariate marginal distributions to form a joint distribution function (Trivedi & Zimmer 2005). Consider two random variables x and y with associated univariate distribution functions $F_x(x)$ and $F_y(y)$. Sklar's (1959) theorem establishes that there exists a copula $C(\,\cdot,\ \cdot\,;\theta)$ such that a bivariate joint distribution is defined for all x and y in the extended real line as

$$
F_{xy}(x, y) = C(F_x(x), F_y(y); \theta) \quad (4)
$$

where the association between the two marginal distributions is represented by the association parameter, θ. This result is remarkable because it shows we can construct a new bivariate distribution based on univariate marginal distributions of known form. As long as the univariate marginal distributions are known, an appropriate choice of copula function C in (4) enables one to represent the unknown bivariate distribution.

Now, the second component of the likelihood function, the probability that disputants reach an agreement at time t_{ci} in conflict and have not yet resumed fighting at least until t_{pi}^0 is obtained by using the Bayes' rule and copula function

$$\begin{aligned}
\Pr(T_{ci} = t_{ci} \cap T_{pi} > t_{pi}^0) &= \Pr(T_{pi} > t_{pi}^0 | T_{ci} = t_{ci}) \times \Pr(T_{ci} = t_{ci}) \\
&= \Pr(T_{pi} > t_{pi}^0 | T_{ci} = t_{ci}) \times f_c(t_{ci}) \\
&= \frac{\partial C\left\{F_p(t_{pi}^0), F_c(t_{ci}); \theta\right\}}{\partial F_c(t_{ci})} \times f_c(t_{ci}).
\end{aligned}$$

Similarly, the third component of (3), the probability that disputants reach an agreement at time t_{ci} in conflict and then resumed fighting at t_{pi} is obtained as

$$\begin{aligned}
\Pr(T_{ci} = t_{ci} \cap T_{pi} = t_{pi}) &= f_{cp}(t_{ci}, t_{pi}) \\
&= \frac{\partial^2 C\left\{F_c(t_{ci}), F_p(t_{pi}); \theta\right\}}{\partial F_c(t_{ci}) \partial F_p(t_{pi})}.
\end{aligned}$$

To complete the derivation, the last step is to choose a particular copula function for $C(,;\theta)$. There are a number of different copula functions that can be used to construct a multivariate distribution from univariate marginals (Trivedi & Zimmer 2005), but some copulas are more flexible than others in that they can accommodate greater range of dependency between the marginals. In this application, I use the Gaussian copula, one of the most flexible copula functions that can accommodate both positive and negative dependency. It has the following form

$$C(u, v; \theta) = \int_{-\infty}^{\Phi^{-1}(u)} \int_{-\infty}^{\Phi^{-1}(v)} \frac{1}{2\pi(1-\theta^2)^{1/2}} \exp\left[\frac{-(s^2 - 2\theta st + t^2)}{2(1-\theta^2)}\right] dsdt$$

where $\Phi^{-1}()$ is the Gaussian quantile function, $-1 < \theta < 1$ is the association parameter, and $u = F_x(x)$ and $v = F_y(y)$ for random variables x and y. The Gaussian copula has a number of desirable characteristics. First, it allows for independence as a special case ($\theta = 0$). We can thus test the existence of interdependence between the two processes by testing whether θ is different from 0. Second, the Gaussian copula is *comprehensive* in that as θ approaches the lower (upper) bound of its permissible range, the copula approaches the theoretical lower (upper) bound. [23] This is not true with other copulas that have been utilized to address selection bias in political science. For example, the estimator proposed by Sartori (2003) forces one to *assume* either one of the theoretical bounds as representing the true data generating process. The consequence of this is not only that we are unable to test the existence of interdependence but also that, depending on the assumption made about the direction of the dependency, we make completely opposite inferences about the effects of explanatory variables on outcomes. The copula function utilized in Boehmke, Morey & Shannon (2006) can accommodate both positive and negative dependency and allows for testing the direction of dependency, but the permissible range is limited to $\theta \in (-0.25, 0.25)$.

[23] The upper and lower theoretical bounds of a joint distribution, called Fréchet bounds, F^- and F^+, are defined as $F^-(u,v) = \max[0, u+v-1]$ and $F^+(u,v) = \min[u,v]$.

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Table 1: Maximum Likelihood Estimates from the Interdependent Duration Model

	Conflict duration	*Peace duration*
IO Membership	-0.70^{**}	-1.10^{**}
	(0.02)	(0.10)
IO Membership $\times$ **log(time)**	0.18^{**}	0.18^{**}
	(0.007)	(0.02)
Capability Balance	0.29^{**}	0.60
	(0.11)	(0.57)
Contiguity	-0.32^{**}	-0.46^{**}
	(0.10)	(0.17)
Major Power	-0.41^{**}	-0.02
	(0.10)	(0.20)
Joint Democracy	-0.18	0.46
	(0.25)	(0.48)
Year of Initiation	-0.01^{**}	-0.01^{*}
	(0.002)	(0.004)
Constant	5.33^{**}	9.02^{**}
	(0.13)	(0.49)
Weibull shape parameter (σ) [(a)]	1.23^{**}	1.21^{**}
	(0.04)	(0.08)
Correlation parameter (θ) [(b)]		-0.21^{**}
		(0.02)
Number of observations	435	427
Number of uncensored obs.	427	168
Time-varying observations	675	7706
Total number of obs.		7954

(*Notes.* As coefficients are shown in the accelerated failure time metric, positive coefficient estimates are associated with longer duration and negative ones are associated with shorter duration. Standard errors in parentheses. Significance levels: $^{*}: 5\%$ $^{**}: 1\%$, two-tailed tests).

[(a)] The Weibull shape parameters (σ) only take positive values. The null hypothesis in testing the significance of σ is $\sigma = 1$.

[(b)] The correlation parameter (θ) takes values between -1 and $+1$.

Table 2: Robustness checks

	(0)	(1)	(2)	(3)	(4)
Conflict duration					
IO Membership	-0.70^{**} (0.02)	-0.66^{**} (0.02)	-0.97^{**} (0.03)	-0.75^{**} (0.01)	-0.67^{**} (0.01)
IO Membership × log(time)	0.18^{**} (0.007)	0.15^{**} (0.007)	0.21^{**} (0.009)	0.18^{**} (0.005)	0.17^{**} (0.006)
Capability Balance	0.29^{**} (0.11)	0.08 (0.28)	0.07 (0.05)	0.34^{**} (0.02)	0.10^{**} (0.004)
Contiguity	-0.32^{**} (0.10)	-0.34^{**} (0.09)	-0.14 (0.08)	-0.37^{**} (0.06)	-0.20^{**} (0.06)
Major Power	-0.41^{**} (0.10)	-0.40^{**} (0.10)	-0.43^{**} (0.10)	-0.39^{**} (0.06)	-0.14^{**} (0.01)
Joint Democracy	-0.18 (0.25)	0.15 (0.23)	0.75^{**} (0.02)	-0.09 (0.11)	-0.41 (0.25)
Year of Initiation	-0.01^{**} (0.002)	-0.01^{**} (0.002)	-0.01^{**} (0.002)	-0.01^{**} (0.001)	-0.01^{**} (0.001)
Constant	5.33^{**} (0.13)	5.66^{**} (0.25)	5.55^{**} (0.07)	5.21^{**} (0.08)	5.37^{**} (0.06)
Weibull shape parameter (σ)	1.23^{**} (0.04)	1.26^{**} (0.05)	1.49^{**} (0.06)	1.19^{**} (0.02)	1.20^{**} (0.02)
Peace duration					
IO Membership	-1.10^{**} (0.10)	-0.71^{**} (0.19)	-1.44^{**} (0.20)	-1.16^{**} (0.07)	-1.05^{**} (0.11)
IO Membership × log(time)	0.18^{**} (0.02)	0.11^{**} (0.03)	0.22^{**} (0.03)	0.19^{**} (0.01)	0.18^{**} (0.02)
Capability Balance	0.60 (0.57)	0.21 (0.73)	0.43 (0.81)	0.57 (0.42)	0.14 (0.66)
Contiguity	-0.46^{**} (0.17)	-1.017^{**} (0.25)	-0.54^{**} (0.25)	-0.51^{**} (0.13)	-0.43^{*} (0.18)
Major Power	-0.02 (0.20)	-0.48 (0.27)	-0.07 (0.29)	-0.18 (0.15)	-0.22 (0.23)
Joint Democracy	0.46 (0.48)	0.70 (0.54)	0.35 (0.73)	-0.01 (0.26)	0.41 (0.49)
Year of Initiation	-0.01^{**} (0.004)	-0.001^{*} (0.005)	-0.01 (0.006)	-0.004 (0.03)	-0.02^{**} (0.005)
Constant	9.02^{**} (0.49)	10.15^{**} (0.49)	9.58^{**} (0.71)	8.61^{**} (0.36)	9.55^{**} (0.59)
Weibull shape parameter (σ)	1.21^{**} (0.08)	0.79^{**} (0.08)	1.20^{*} (0.11)	1.20^{**} (0.06)	1.29^{**} (0.10)
Correlation parameter (θ)	-0.21^{**} (0.02)	0 (assumed)	-0.18^{**} (0.02)	-0.19^{**} (0.01)	-0.23^{**} (0.02)
N: Conflict duration	435	435	297	595	336
N: Peace duration	427	427	289	586	336
N: Total	7954	7954	5907	9483	5959
Log likelihood	-4188	-4289	-2618	-6231	-3270

(0) The baseline model, as shown in Table 1.
(1) A naive model, where the two durations are assumed independent.
(2) Analyzing conflicts that involve serious clashes.
(3) Including non-violent conflicts.
(4) Post-1945 conflicts only.

Table 3: Interdependent Duration Analysis with Selection

	1st Stage *(Peace before conflict)*	*2nd Stage* *(Conflict duration)*	*3rd Stage* *(Peace after conflict)*
IO Membership	-0.45	-0.64^{**}	-0.77^{**}
	(0.28)	(0.06)	(0.18)
IO Membership × log(time)	0.05	0.11^{**}	0.12^{**}
	(0.04)	(0.02)	(0.03)
Capability Balance	2.14^{*}	2.28^{**}	-0.94
	(1.09)	(0.59)	(0.85)
Contiguity	-1.57^{**}	-0.61^{**}	0.09
	(0.49)	(0.23)	(0.40)
Major Power	-0.51	-0.58^{**}	-0.01
	(0.40)	(0.19)	(0.33)
Joint Democracy	1.84^{**}	0.89^{*}	$2.09^{\dagger}$
	(0.57)	(0.35)	(1.09)
Year of Initiation	0.004	-0.01^{**}	0.001
	(0.007)	(0.004)	(0.008)
Constant	8.56^{**}	4.58^{**}	8.78^{**}
	(0.96)	(0.43)	(0.72)
Weibull shape parameter (σ)	0.62^{**}	1.26^{**}	1.32^{**}
	(0.06)	(0.12)	(0.15)
Correlation parameter (θ) [(a)]		0.45^{**}	-0.15^{**}
		(0.14)	(0.04)
Number of Observations	249	131	113
Number of Uncensored Obs.	131	113	59
Time-varying Observations	3097	197	951

(*Notes.* Standard errors in parentheses. Significance levels: † : 10% * : 5% ** : 1%, two-tailed tests).

[(a)] The correlation parameter under the second column (0.45) measures the correlation between residual durations from the 1st and 2nd stages, whereas that under the third column (-0.15) measures the correlation between 2nd and 3rd stages.