The Strength of Cease-fire Agreements and the Duration of Postwar Peace*

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Abstract

Do stronger cease-fire agreements keep peace longer after war? Although there are theoretical reasons to expect that stronger agreements promote durable peace, the extant empirical research provides mixed support for this expectation. This research note reexamines this argument empirically, addressing two inferential problems that have not been fully addressed in the past studies. First, since the strength of cease-fire agreements is endogenous to the baseline prospect for peace, I develop a new statistical model that jointly estimates agreement strength and peace duration in a unified framework. Second, I allow the effect of agreement strength to vary over time. This is important because agreement strength may matter little right after the war, for there exists a rough consensus among the ex-belligerents about the likely outcome of a next war. As time passes, however, there will be a greater chance that some exogenous shocks distort this consensus that had facilitated war termination in the first place, and that is when the effect of agreement strength will start to show. Analyzing the duration of postwar peace from 1914 to 2001, I demonstrate that stronger cease-fire agreements indeed stabilize peace.

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Introduction

Do stronger cease-fire agreements keep peace longer after war? When ex-belligerents sign an agreement that incorporates institutional mechanisms designed to alter the strategic environment, are they able to maintain more durable peace? Scholars of international relations have argued that international agreements can change states' behaviors even when there is no central authority that enforces an agreement. Studies have demonstrated that international agreements can indeed constrain states' behavior and promote cooperation in the absence of direct enforcement in wide range of issue areas, such as international security,[1] human rights protection,[2] and monetary affairs.[3] In line with these studies, Fortna argues that stronger cease-fire agreements, characterized by mechanisms such as confidence building measures, dispute settlement procedures, peacekeeping forces, etc., are able to help ex-belligerents overcome the obstacles to peace, hence prolonging the duration of peace after war.[4]

However, although we have compelling theoretical reasons to believe that stronger cease-fire agreements promote durable peace, the extant empirical findings on this relationship are rather mixed. On the one hand, Fortna finds empirical support for her argument using data on duration of peace after interstate wars for the period between 1946 and 1998.[5] On the other hand, subsequent studies have provided little support for this relationship. Analyzing the data from the same period, Werner and Yuen find that agreement strength has little or no effect, depending on model specifications.[6] Moreover, a more recent study by Lo, Hashimoto, and Reiter also finds no effect using an expanded data set that covers the interstate wars from 1914 to 2001.[7]

This research note reevaluates this argument empirically and demonstrates that agreement strength indeed prolongs postwar peace duration. In doing so, I address

1 See Leeds 2003; Mattes 2008; Morrow 2007 for example.
2 See Hathaway 2002; Hill 2010.
3 See Simmons 2000.
4 Fortna 2003 and 2004.
5 Fortna 2003 and 2004.
6 Werner and Yuen 2005.
7 Lo, Hashimoto, and Reiter 2008.

two inferential problems that have plagued previous attempts to uncover the true effect of agreement strength. The first challenge is that agreement strength is possibly endogenous to the prospect for peace: ex-belligerents will seek to reach stronger cease-fire agreements when they have more fragile prospect for peace. This means that, if we fail to correct for this endogeneity, our estimate of the peace-enhancing effect of agreements will be negatively biased. To address this issue, I develop an empirical strategy that jointly estimates agreement strength and peace duration while controlling for the effect of unobservable factors that influence the both processes.

The second inferential problem is the assumption that agreement strength has a time-constant effect on peace duration. As I argue in greater details below, the effect of agreement strength will be likely to grow stronger over time as the underlying conditions that have led to the war-ending agreement will change over time. I relax this assumption by introducing an interaction term between agreement strength and the analysis time. The empirical analysis shows that, if we address these two problems, agreement strength can be shown to prolong the duration of peace after wars.

The remainder of this note is organized as follows. In the next section, I discuss theoretical reasons as to why we should expect stronger cease-fire agreements to foster longer peace after war. I then identify two inferential problems that have hampered previous empirical efforts to correctly identify the effect of agreement strength. The following sections introduce the research design and present the empirical findings. The final section concludes by summarizing the implications for other studies and policy makers.

Why Stronger Agreements Promote Peace

Why does war occur? This has been one of the most important questions that concerns scholars of international relations for a long time. Thanks to the development of bargaining theories, the past two decades have seen a significant development in the scholarly efforts to answer this question. The bargaining model of war attributes the causes of war to a failure of pre-war bargaining between disputants. Given that war is costly and that most wars end in an agreement, under broad conditions there will exist a bargain (= a division of the disputed good) that can make both sides better off than fighting. However, pre-war bargaining can break down in costly fighting when 1) disputants have uncertainties about the underlying balance of power (and hence about the outcome of war), or 2) disputants cannot commit to following through on an agreement in the absence of central enforcement.[8]

One of the strengths of the bargaining approach to war is that it can not only explain the outbreak of war but also offer insight into the duration, termination, and recurrence of war in a coherent framework. If we argue that a certain factor is a cause of a war, logical consistency requires that, for the war to end that factor must be eliminated.[9] For example, if states fight because both sides are uncertain about the other sides' military power, then the war should end when both sides have learned enough about the other sides' strength such that neither side has incentive to continue fighting. This means that, at war's end, the disputants should be 1) in agreement about the underlying power balance and 2) capable of committing to following through on a cease-fire; otherwise, the disputants would keep fighting. This argument further implies that postwar peace may break down in another war when the underlying conditions that have led to the termination of fighting change in the future.[10]

However, war recurrence is not always inevitable, even when such a change in underlying conditions occurs and the war-ending settlement becomes obsolete. One viable way for ex-belligerents to maintain postwar peace in the face of changing conditions is to devise strong cease-fire agreements that incorporate institutional mechanisms to help them resolve bargaining problems that may arise in the future. For example, some cease-fire agreements incorporate institutional devices such as dispute settlement mechanisms or confidence building measures that will provide the ex-belligerents with information about the actions and intentions of the other side, thus reducing un-

[8] Fearon 1995; Powell 2006.
[9] Reiter 2009; Wagner 2000.
[10] Werner 1999.

certainties that may precipitate violence. Those agreements that are characterized by the presence of third-party involvement or peacekeeping force will enhance the ex-belligerents' ability to commit to compliance with the cease-fire by raising the reputational cost or audience cost of noncompliance.[11] In addition, Fortna argues that stronger cease-fire agreements can also contribute to longer peace by controlling accidents.[12] Those agreements that incorporate arms control provisions, confidence building measures, demilitarized zones, or specify treaty obligations in more precise wordings, will be able to prevent accidents or involuntary defection from happening, or prevent such accidents from escalating into more violent conflicts.

Given that it would be costly for disputants to devise, negotiate, and implement these mechanisms, they would not bother signing stronger agreements if they did not expect agreement strength to make a difference. Nevertheless, previous studies have been unable to find a robust relationship between agreement strength and the duration of postwar peace. Whereas Fortna's initial statistical analyses that cover the period from 1946 to 1998 lend support for this expectation, subsequent studies have found little support. Analyzing the same time period, Werner and Yuen find that agreement strength matters little once they control for some conditions under which bargaining obstacles are particularly likely to arise.[13] They conclude that institutional devices incorporated in cease-fire agreements may not be strong enough when the war-ending settlement becomes obsolete as a result of changes in the strategic environment. Moreover, a more recent study by Lo, Hashimoto, and Reiter that covers a longer time period finds that agreement strength has no discernible effect on postwar peace duration.[14]

I argue that previous studies have failed to find strong support for this relationship because they do not offer a fair test of the theoretical argument linking agreement strength and post-war peace duration. In what follows, I identify two inferential problems that have plagued previous research, and then propose a solution to overcome these problems.

11 Mattes 2008; Mattes and Vonnahme 2010.
12 Fortna 2003, 2004.
13 Werner and Yuen 2005.
14 Lo, Hashimoto, and Reiter 2008.

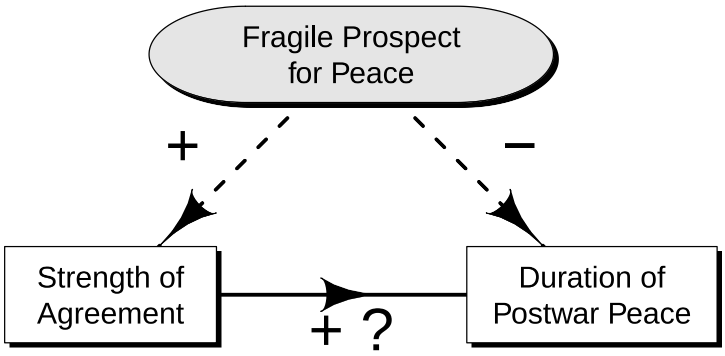

Figure 1: Endogeneity of agreement strength

Notes. This figure graphically illustrates the endogeneity problem. Observable variables (Strength of Agreement and Duration of Postwar Peace) are shown in white squares, and an unobservable variable (Fragile Prospect for Peace) is shown in a gray ellipse.

Two Problems with Existing Studies

The first inferential problem is endogeneity of agreement strength. State leaders carefully design international agreements and institutions to advance their interests,[15] which means that we need to take into account resulting self-selection and endogeneity when assessing their effect on leaders' behaviors.[16] As Fortna demonstrates in her study of agreement strength, disputants tend to sign stronger agreements when they expect the postwar peace to be more unstable.[17] Put differently, agreement strength will be positively correlated with a fragile prospect for peace, as illustrated in the left-had side of Figure 1. As fragile prospect for peace will be negatively correlated with the duration of postwar peace, a failure to control for the effect of fragile prospect for peace would result in a negative omitted variable bias in our assessment of the relationship between agreement strength and peace duration. In other words, the endogeneity of agreement strength can make it look that agreement strength is unrelated to postwar peace duration even when it has a positive effect in truth.

Recognizing this endogeneity, researchers have sought to alleviate the problem by controlling for *observable* proxy variables that are supposedly correlated with the un-

15 Koremenos, Lipson, and Snidal 2001.
16 Downs, Rocke, and Barsoom 1996.
17 Fortna 2004.

observable fragile prospect for peace. However, disputants' expectation (i.e., Fragile Prospect for Peace) is largely unobservable to the researcher, and so it is difficult if not impossible to measure and control for it. Even after controlling for several observable proxies of it, there may still exist some unobservable factors that are negatively correlated with the duration of peace and positively influencing the strength of agreement. If we fail to control for such residual unobservables, our empirical results will still be biased against finding the expected positive effect of stronger agreements on peace duration. In order to correctly estimate the true effect of agreement strength on peace duration in the face of endogeneity, we need an empirical approach that corrects for the influence of *unobservable* factors. The next section introduces an original statistical model that accounts for the correlation between the error terms governing duration and agreement strength specifically developed for this study.

The second problem is that the existing studies on postwar peace duration have assumed that the effect of cease-fire treaties is constant over time, although it will probably vary over time. Mattes argues that, from the perspective of the bargaining theory of war, the constraining effect of international agreements and institutions will start to show only after the strategic conditions at the time of agreement formation change over time.[18] Since states self-select into an international agreement that is consistent with their preference, states would find it easy to comply with the agreement as long as the strategic environment at the time of agreement formation remains unchanged. In such situations, the observed state compliance cannot be attributed to the constraining effect of an agreement but is due simply to prior self-selection.[19] If, however, states still comply with the agreement even after the strategic environments change over time, we can say that it is (at least partially) because of the constraining effect of agreement. I will apply this argument to understanding the effect of agreement strength on the stability of peace after war.

Immediately after a war ends, there will exist a rough consensus among the ex-belligerents about the distribution of power between them, and hence about the likely

[18] Mattes 2008.
[19] Downs, Rocke, and Barsoom 1996.

outcome of a next war should one happen. The bargaining theory of war discussed above maintains that this convergence of expectation is the very reason why the disputants stop fighting in the first place. Therefore, right after the war ends, there would be few sources of bargaining problems, which means that the agreement strength would not make much of a difference. However, as time passes, there will be a greater chance that some exogenous shocks distort this consensus, creating bargaining problems. It may be the case that some new issues to fight over arise between the ex-belligerents, or a more hawkish leader comes into power in one country who sees greater benefit in challenging the previously agreed status quo, or one country's military power grows faster than that of the other country, which creates dynamic commitment problems. These are the situations in which the agreement strength will make a greater difference. If we were to assume that the effect of agreement strength is time-constant, we would underestimate the true effect of agreement strength. I deal with this problem by introducing an interaction term between time and the strength of agreement, thereby allowing the effect of agreement strength to vary over time.

Research Design

I now turn to the empirical analysis of agreement strength and postwar peace duration. To maintain consistency with the previous studies, I utilize the data set originally developed by Fortna and later expanded by Werner and Yuen and Lo, Hashimoto, and Reiter.[20] The basic unit of observation is a war-dyad, a pair of countries that fought an interstate war against each other. The outcome variable is the duration of peace measured in days following interstate wars that began between 1914 and 2001. There are 52 wars in the data set, involving 186 war-dyads. Among the 186 war-dayds, 54 failures of peace are observed. For the remaining 132 observations, post-war peace did not break down as of the end of 2001 and thus they are treated as right-censored. Since many of the covariates change their values by year, the 186 war-dyad observations are divided into multiple annual observations, yielding 6,368 war-dyad-years.

[20] Fortna 2003; Werner and Yuen 2005; Lo, Hashimoto, and Reiter 2008.

The key explanatory variable is the strength of cease-fire agreements reached by the members of war-dyads. This variable is a composite index measuring the presence and strength of various mechanisms intended to stabilize peace after wars, such as demilitarized zone, arms control measures, peacekeeping force, etc. There are ten constitutive variables, each of which takes equidistant values between 0 and 1.[21] I transform this index into an integer-valued variable that ranges between 0 and 60, without changing the substantive interpretation of the original index developed by Fortna.[22] Fortna initially developed the data for the period between 1945 and 1998, and the data were subsequently expanded for the period between 1919 and 2001 by Lo, Hashimoto, and Reiter using Fortna's codebook. [23] Overall, the dataset contains information about 136 peace agreements, including cease-fire agreements and follow-up agreements.

Statistical Model

To address the endogeneity problem, I need a statistical model that explains peace duration as a function of an endogenous regressor (i.e., agreement index) and some covariates, while controlling for the correlation between the error terms governing the two. Since there is no "off-the-shelf" estimator readily available for my purpose, I develop a full-information maximum likelihood estimator that jointly estimates the two processes. The model is an extension of the standard instrumental regression procedure utilized to address endogeneity, and shares many features with other multi-equation statistical models.

For each war-dyad $i = 1, \cdots, n$, we define a latent discrete random variable, s_i^*, to represent the belligerents' propensity to reach stronger agreements.[24] As agreement

21 Table A.1 in the Appendix describes the measurement of the ten mechanisms that comprise this index. Five of the constitutive variables are binary (0 or 1), four of them take three values $(0, \frac{1}{2}, 1)$, and one of them takes four values $(0, \frac{1}{3}, \frac{2}{3}, 1)$. Fortna simply sums up these 10 variables to create a simple index that ranges between 0 and 10.

22 Specifically, I recoded this index by multiplying it by 6. This means that binary constitutive variables $a \in \{0, 1\}$ are recoded into $a' \in \{0, 6\}$, three-valued constitutive variables $b \in \{0, \frac{1}{2}, 1\}$ are recoded into $b' \in \{0, 3, 6\}$, and four-valued constitutive variables $c \in \{0, \frac{1}{3}, \frac{2}{3}, 1\}$ are recoded into $c' \in \{0, 2, 4, 6\}$, and so the resulting index is recoded to take integer values between 0 and 60. This linear transformation is done to facilitate estimation and involves no loss of information.

23 Fortna 2003, 2004; Lo, Hashimoto, and Reiter 2008.

24 Since agreement strength can change values over time, a given war-dyad can have multiple values

strength only takes on non-negative integers, we can model it with a Poisson distribution with the cumulative distribution function $F_s(s_i) \equiv \Pr(s_i^* < s_i)$ where s_i is a realized value of s_i^*. In addition, for each war-dyad i we define another latent random variable, t_i^*, to represent duration of postwar peace. To capture the effect of s_i on t_i^*, we allow the duration of postwar peace to be conditioned on s_i and a set of covariates as follows:

$$\log(t_i^*) = \beta s_i + \boldsymbol{x_i}\boldsymbol{\gamma} + \alpha^{-1}\epsilon_i \tag{1}$$

where β is the effect of agreement index on duration, $\boldsymbol{x_i}$ is a vector of covariates, $\boldsymbol{\gamma}$ is a vector of coefficients, and ϵ_i is a stochastic disturbance term scaled by α. If we assume ϵ_i follows a type-1 extreme-value distribution (also known as log-Weibull distribution), we obtain a Weibull duration model.[25]

However, univariate estimation of (1) would not give us an unbiased estimate of β because some unobservable factors may influence both s_i and t_i simultaneously. For example, suppose belligerents tend to sign stronger agreements when they have a more fragile prospect for peace. This implies that the unobservable component of s_i will be negatively correlated with ϵ_i, which is the unobservable component of t_i. Then, when s_i takes an unusually high value, ϵ_i will on average take an unusually low value. Unless we account for the correlation between these two unobservable components, we will underestimate β, because part of the effect of ϵ_i on t_i^* will be picked up by β. In other words, we will incorrectly attribute to β the effect of any part of ϵ that is correlated with the unobservable component of s_i.

To obtain an unbiased estimate of β, we thus need to construct a statistical model that jointly estimates s_i^* and t_i^* while controlling for the correlation between unobservables that influence both. The likelihood function of such a statistical model takes the

of agreement strength. However, to simply the exposition of the model, I only discuss a simplified version of the model that assumes away time-varying covariates (i.e., the model only allows for one value of s_i^* per one war-dyad). Appendix A presents the derivation of the full model that incorporates time-varying covariates.

25 I chose Weibull distribution for illustrative purposes here. In the empirical application that follows, I estimate models with several different distributions and choose the best fit model based on fit statistics. I will revisit this point later.

following form:

$$\mathcal{L} = \prod_{i=1}^{n} \Pr(t_i^* = t_i \cap s_i^* = s_i)^{r_i=0} \Pr(t_i^* > t_i^0 \cap s_i^* = s_i)^{r_i=1}, \tag{2}$$

where t_i is the observed duration for war-dyad i until the time of war recurrence, t_i^0 is the observed duration for war-dyad i until the time of right-censoring, and r_i is a right-censoring indicator coded 1 for censored and 0 for uncensored observations. Thus the first component of (2) calculates the likelihood contribution from uncensored observations, whereas the latter part represents that from censored ones.

To fully specify this likelihood function, we need to characterize two joint distributions of t^* and s^* that appear in (2). I utilize a copula function to characterize these bivariate distributions where one marginal is Weibull and the other marginal is Poisson. A copula is a function that binds together multiple univariate distributions of known form to produce a joint probability distribution.[26] For two random variables x^* and y^* with associated univariate distribution functions $F_x(\cdot)$ and $F_y(\cdot)$, the bivariate distribution can be defined as $F_{xy}(x, y) \equiv \Pr(x^* < x \cap y^* < y) = C\{F_x(x), F_y(y); \theta\}$, where $C(\cdot, \cdot; \theta)$ is a bivariate copula function and θ parameterizes the degree of association between x and y. What is remarkable about copula functions is that it enables us to derive a joint statistical model of multiple processes of any form as long as we know each of the univariate marginal distributions.[27]

The first component of (2), the joint probability that disputants reach an agreement with strength s_i and then war recurs after duration t_i is obtained by applying the Bayes'

26 For an introduction to copula functions, see Trivedi and Zimmer 2005.

27 Recent years have seen a handful of political science applications of copula-based joint modeling. For example, Chiba, Martin, and Stevenson (2015) propose a duration model with sample selection where the selection process is multichotomous choice; Chiba, Metternich, and Ward (2015) propose models of multiple consecutive duration processes; and Fukumoto (2015) proposes joint models of duration and an ordered outcome.

rule and taking the derivatives of the joint distributions, as follows:

$$\begin{aligned}
\Pr(t_i^* = t_i \cap s_i^* = s_i) &= \Pr(t_i^* = t_i \cap s_i^* < s_i) - \Pr(t_i^* = t_i \cap s_i^* < s_i - 1) \\
&= \Pr(s_i^* < s_i | t_i^* = t_i) \times f_t(t_i) - \Pr(s_i^* < s_i - 1 | t_i^* = t_i) \times f_t(t_i) \\
&= \left[\frac{\partial C\left\{F_s(s_i), F_t(t_i); \theta\right\}}{\partial F_t(t_i)} - \frac{\partial C\left\{F_s(s_i - 1), F_t(t_i); \theta\right\}}{\partial F_t(t_i)} \right] \times f_t(t_i) \\
&= \left[C_{u|v}\left\{F_s(s_i), F_t(t_i); \theta\right\} - C_{u|v}\left\{F_s(s_i - 1), F_t(t_i); \theta\right\} \right] \times f_t(t_i) \quad (3)
\end{aligned}$$

where $f_t(\cdot)$ and $F_t(\cdot)$ are the univariate density and distribution functions for t, and $C_{u|v}(u, v; \theta)$ is the conditional copula that represents $\Pr(u^* < u \mid v^* = v)$. Note that, as the agreement strength is a discrete variable, $\Pr(s_i^* = s_i)$ is obtained as $\Pr(s_i^* < s_i) - \Pr(s_i^* < s_i - 1)$. Similarly, the second component of the likelihood function, the joint probability that disputants reach an agreement with strength s_i and then peace has lasted for length t_i^0 until the observation is right-censored, is obtained as:

$$\begin{aligned}
\Pr(t_i^* > t_i^0 \cap s_i^* = s_i) &= \Pr(t_i^* > t_i^0 \cap s_i^* < s_i) - \Pr(t_i^* > t_i^0 \cap s_i^* < s_i - 1) \\
&= \Pr(s_i^* < s_i) - \Pr(t_i^* < t_i^0 \cap s_i^* < s_i) - \left\{\Pr(s_i^* < s_i - 1) - \Pr(t_i^* < t_i^0 \cap s_i^* < s_i - 1)\right\} \\
&= F_s(s_i) - F_s(s_i - 1) - C\left\{F_s(s_i), F_t(t_i^0); \theta\right\} + C\left\{F_s(s_i - 1), F_t(t_i^0); \theta\right\}. \quad (4)
\end{aligned}$$

The final step to complete the derivation of the likelihood function is to choose a particular copula function, $C(,;\theta)$. I use the Gaussian copula, which takes the following form:

$$\begin{aligned}
C(u, v; \theta) &= \Phi_2\left\{\Phi^{-1}(u), \Phi^{-1}(v); \theta\right\} \\
C_{u|v}(u|v; \theta) &= \Phi\left\{\frac{\Phi^{-1}(u) + \theta\Phi^{-1}(v)}{\sqrt{1 - \theta^2}}\right\}, \quad (5)
\end{aligned}$$

where $u = F_x(x)$, $v = F_y(y)$, $\Phi(\cdot)$ is the distribution function and $\Phi^{-1}(\cdot)$ is the quantile function of the standard Normal distribution, $\Phi_2(\cdot)$ is the distribution function of the standard bivariate Normal distribution, and θ is the association parameter that ranges between -1 (perfect negative correlation) and 1 (perfect positive correlation). Gaussian copula is one of the most flexible copula functions that can accommodate both positive and negative dependency to the maximum extent.

Covariates and Identification

To identify the model without relying solely on distributional assumptions about nonlinearity, we need two sets of covariates, one explaining both agreement strength s and duration t and the other set explaining only s. In other words, we need to identify some correlates (often called instruments) of agreement strength that are not *directly* related to peace duration other than through their correlations with agreement strength. As for the first set of variables included in both equations, I rely on previous studies of agreement strength and peace duration.[28] Some of the covariates are time-varying: `Relative Capabilities` is military capability (CINC) score of the stronger state divided by the sum of capability scores of the dyad based on Correlates of War Project (COW); and `Fluctuation in Capabilities` is year-to-year fluctuation in `Relative Capabilities`; `One Democracy` is a dummy variable equal to 1 if only one member of the dyad is a democracy (Polity score equal to or greater than 5) and 0 otherwise; and `Joint Democracy` is another dummy variable equal to 1 if both are a democracy. Other covariates are time-invariant and measured at war's end.[29]

As for the instruments that are included only in the agreement strength equation, I prepare three variables that capture changes in the strategic environment surrounding the ex-belligerents. The idea is that such changes can become the sources of bargaining failures, which, in turn, should give the ex-belligerents incentives to strengthen peace agreements. At the same time, environmental changes are not *directly* related to

[28] All the variables are taken directly from Lo et al's data set, which is based on Fortna's and Werner and Yuen's coding rules. Fortna 2003, 2004; Werner and Yuen 2005; Lo, Hashimoto, and Reiter 2008.

[29] Time-invariant covariates include: `Foreign-imposed Regime Change` (equal to 1 if the war winner imposed new government on the loser, 0 otherwise), `Military Tie` (equal to 1 if the war ended in a military tie, 0 otherwise), `Casualty` (natural log of battle death for the dyad), `Existence at Stake` (equal to 1 if the war involved threat to existence, 0 otherwise), `Multilateral War` (equal to 1 for multilateral wars, 0 for bilateral wars), `Battle Consistency` (a measure of the amount of information exchanged in the battlefield), `Interrupted War` (equal to 1 if the war was interrupted by a third-party, 0 otherwise), `Contiguity` (equal to 1 if the disputants are geographically contiguous, 0 otherwise), and `Conflict History` (number of prior militarized interstate disputes divided by number of years both states are in the interstate system since 1816 at time the war started). See Lo, Hashimoto, and Reiter 2008 for detailed coding rules.

peace duration but related *only through* changes in agreement strength, as long as the ex-belligerents are responsive to such changes. Therefore, this additional variation in agreement strength due to environmental changes will enable us to identify the true effect of agreement strength in the face of endogeneity. I measure changes in the strategic environment by focusing on military balance and regime type.[30] `Change in Relative Capabilities` is the absolute value of the difference between `Relative Capabilities` measured at war's end and that measured in the observation year. Since a large change in the balance of power can lead to dynamic commitment problems as well as uncertainties about the balance of power, I expect that this variable is positively associated with agreement strength. Second, `One Democracy Same` and `Joint Democracy Same` are dummy variables, each coded as 1 if the respective indicator of regime type in the observation year takes the same value as that at war's end, and as 0 otherwise. Since changes in regime type can be a source of bargaining obstacles, I expect that these variables are negatively associated with agreement strength.

Estimation Results

Table 1 reports the maximum likelihood estimates of model parameters. Among the four models presented in the table, the main focus of this study is Model (4) since this is the one that addresses both of the two inferential problems discussed above. I report the results from three other models to facilitate a comparison between my findings with those we would obtain if failed to address one or both of the problems. Model (1) presents a simple univariate analysis of postwar peace duration that ignores both of the two problems. Model (2) ignores the endogeneity, but it relaxes the time-invariant effect assumption by including an interaction time between agreement strength and time. Model (3) corrects for the endogeneity, but it omits the interaction term. Model (4) corrects for the endogeneity and includes the interaction term to relax the time-invariant assumption. Since Model (4) yields the best (i.e., lowest) AIC score, we can

30 For studies that adopt a similar approach, see Leeds 2003; Mattes 2008.

say that Model (4) fits the data better than the other three models.[31]

Table 1: Models of agreement strength and postwar peace duration

	Model (1)	Model (2)	Model (3)		Model (4)	
	Peace Duration [Weibull]	Peace Duration [Weibull]	Peace Duration [Weibull]	Agreement Strength [Poisson]	Peace Duration [Weibull]	Agreement Strength [Poisson]
Agreement Strength	0.010	-0.114^{**}	0.010		-0.074^{***}	
	(0.022)	(0.054)	(0.010)		(0.009)	
Agreement Strength $\times$ log(t)		0.017^{**}			0.012^{***}	
		(0.007)			(0.001)	
Change in Relative Capabilities				0.613^{***}		0.613^{***}
				(0.043)		(0.032)
One Democracy Same				-0.141^{***}		-0.141^{***}
				(0.006)		(0.006)
Joint Democracy Same				0.002		0.002
				(0.014)		(0.013)
Foreign-imposed Regime Change	2.611^{*}	1.990^{*}	2.611^{***}	-0.373^{***}	1.989^{***}	-0.373^{***}
	(1.378)	(1.064)	(0.154)	(0.009)	(0.297)	(0.008)
Military Tie	-1.376^{*}	-0.923	-1.376^{***}	0.173^{***}	-0.930^{***}	0.173^{***}
	(0.721)	(0.585)	(0.225)	(0.009)	(0.162)	(0.007)
Casualty	0.476^{***}	0.342^{***}	0.476^{***}	0.027^{***}	0.338^{***}	0.027^{***}
	(0.137)	(0.115)	(0.004)	(0.001)	(0.0001)	(0.0001)
Existence at Stake	-2.414^{***}	-1.770^{***}	-2.414^{***}	-0.099^{***}	-1.892^{***}	-0.099^{***}
	(0.746)	(0.615)	(0.345)	(0.006)	(0.004)	(0.006)
Multilateral War	0.606	0.502	0.606^{**}	0.075^{***}	0.496^{***}	0.075^{***}
	(0.906)	(0.678)	(0.236)	(0.010)	(0.161)	(0.005)
Battle Consistency	1.103	0.871	1.103^{***}	-0.062^{***}	0.923^{***}	-0.061^{***}
	(0.754)	(0.575)	(0.003)	(0.007)	(0.009)	(0.006)
Interrupted War	-0.371	-0.187	-0.371	0.040^{***}	-0.161	0.040^{***}
	(0.707)	(0.536)	(0.563)	(0.010)	(0.161)	(0.009)
Contiguity	-2.315^{***}	-1.812^{***}	-2.315^{***}	0.092^{***}	-1.815^{***}	0.092^{***}
	(0.768)	(0.629)	(0.306)	(0.006)	(0.161)	(0.006)
Conflict History	-0.822	-0.582	-0.822^{***}	0.047^{***}	-0.605^{***}	0.047^{***}
	(0.859)	(0.652)	(0.006)	(0.014)	(0.001)	(0.0004)
Relative Capabilities	-5.997^{***}	-4.428^{**}	-5.997^{***}	0.494^{***}	-4.414^{***}	0.494^{***}
	(2.090)	(1.738)	(0.005)	(0.022)	(0.0005)	(0.0003)
Fluctuation in Capabilities	-0.829^{*}	-0.578	-0.829^{***}	-0.036^{***}	-0.631^{***}	-0.036^{***}
	(0.452)	(0.358)	(0.056)	(0.012)	(0.0005)	(0.0004)
One Democracy	-0.065	-0.110	-0.065	-0.056^{***}	-0.088	-0.056^{***}
	(0.583)	(0.438)	(0.640)	(0.007)	(0.161)	(0.006)
Joint Democracy	3.130^{**}	2.432^{**}	3.130^{***}	0.123^{***}	2.432^{***}	0.123^{***}
	(1.372)	(1.111)	(0.915)	(0.013)	(0.178)	(0.013)
Constant	13.079^{***}	11.876^{***}	13.079^{***}	2.621^{***}	11.864^{***}	2.621^{***}
	(2.278)	(1.828)	(1.169)	(0.030)	(0.323)	(0.016)
Duration Dependence: $\log(\alpha)$	-0.572^{***}	-0.286	-0.572^{***}		-0.271^{***}	
	(0.124)	(0.192)	(0.006)		(0.0002)	
Correlation: $\tanh^{-1}(\theta)$			-0.067^{***}		-0.473^{***}	
			(0.0002)		(0.0001)	
AIC	59708	59706	59380		58865	

Standard errors in parentheses. 186 subjects, 6368 observations. $^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$

31 To obtain AIC scores for univariate models (i.e., Models (1) and (2)) that are comparable to those for bivariate models (i.e., Models (3) and (4)), I estimated the corresponding bivariate models by assuming a zero correlation between the error terms. Specifically, the parameters for Model (1) are obtained by estimating Model (3) with the assumed zero correlation, and the parameters for Model (2) are obtained by estimating Model (4) with the assumed zero correlation.

In the Peace Duration equations (columns 1, 2, 3, and 5), covariates with a positive coefficient are associated with longer duration since estimates are shown in the accelerated failure time metric. I chose the Weibull specification to model the duration process based on fit statistics. I also estimated these four models with the Log-logistic specification to test if the underlying hazard exhibits a non-monotonous shape, but these models yield inferior AIC scores. The estimated duration dependence parameter $\log(\alpha)$ is negative, suggesting that the underlying risk of war recurrence controlling for the effect of covariates is decreasing over time.[32] The estimated correlation parameter $\tanh^{-1}(\theta)$ is negative in both Models (3) and (4), consistent with the expectation that unobservable factors associated with shorter duration of peace are correlated with unobservable factors associated with stronger agreements.[33] That is, all else equal, disputants tend to sign stronger agreements when they have a fragile prospect for peace. This finding goes against a claim that ex-belligerents sign stronger agreements only when they can easily comply with them and maintain stable peace without a strong agreement anyway,[34] which would imply a positive correlation between the error terms.

In the Agreement Strength equations (columns 4 and 6), covariates with a positive coefficient are associated with stronger agreements since they represent their effects on the mean of a Poisson distribution describing agreement strength. Estimated coefficients are largely in the expected direction and consistent with previous findings. More importantly, two of the three instruments included only in the Agreement Strength equation are highly statistically significant and have the correct signs. Although there is no formal test of the strength of instruments for a non-linear model like this one, this piece of evidence suggests that these variables are collectively strong enough to identify the model.

32 α is the scale parameter for a Weibull distribution, where $\alpha > 1$ implies positive duration dependence (the risk of failure increases over time), $0 < \alpha < 1$ implies negative duration dependence, and $\alpha = 1$ implies no duration dependence. Since α must be strictly positive, I estimate $\log(\alpha)$ instead, which ranges between $-\infty$ and ∞. With this transformation, a positive (negative) estimate of $\log(\alpha)$ implies positive (negative) duration dependence.

33 Since θ is constrained between -1 and 1, I estimate $\tanh^{-1}(\theta)$ (the inverse hyperbolic tangent of θ) instead, which ranges between $-\infty$ and ∞. With this transformation, $\tanh^{-1}(\theta) = -0.067$ implies $\theta \simeq -0.067$ and $\tanh^{-1}(\theta) \simeq -0.473$ implies $\theta = -0.441$.

34 For a compelling argument consistent with such a claim, see Downs, Rocke, and Barsoom 1996.

Let us now turn to the estimated effects of agreement strength implied by these four models. Model (1) yields a positive but statistically insignificant coefficient for Agreement Strength, suggesting that we would mistakenly conclude that agreement strength has no statistically discernible effect on peace duration if we were to ignore the endogeneity and time-varying effect of agreement strength. This is the finding typically reported in previous studies of postwar peace duration.[35] Correcting for the endogeneity alone would not alter this conclusion, as suggested by Model (3). However, if we also allow the effect of agreement strength to vary over time, the agreement strength seems to have statistically significant effect on peace duration, as suggested by Models (2) and (4). Since agreement strength is interacted with the natural log of analysis time in these models, the *implied* coefficient for agreement strength will vary over time.[36]

It is important to note, however, that, in multi-equation, non-linear statistical models such as this model, estimated raw coefficients are not necessarily indicative of the substantive effects of the covariates on the outcome variable of interest. In fact, when the estimated correlation between multiple equations is not zero and a covariate appears in more than one equations, it is often the case that the magnitude, sign, and statistical significance of the substantive effect of the variable are different from its coefficient.[37] Therefore, I illustrate the estimated impact of agreement strength on peace duration by calculating the change in survival probability of peace corresponding to a

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[35] While the Fortna study and my study adopt the Weibull specification, which relies on the assumption that the underlying hazard is either monotonically increasing or monotonically decreasing, Werner and Yuen (2005) and Lo, Hashimoto, and Reiter (2008) both adopt a Cox semi-parametric specification that does not require assumptions about the shape of the hazard. The Cox approach is supposedly more flexible than parametric duration models such as Weibull model, but parametric models are more efficient when the distributional assumptions are met. The choice between different specifications is ultimately an empirical question. As mentioned above, the specification test comparing Weibull and Log-logistic models suggests that there is no evidence of non-monotonic hazard. That said, specification choice does not make much of a difference in this particular application: the results shown in Model (1) are substantively similar to the results from the Cox model reported in previous studies.

[36] More specifically, according to Model (4), the implied coefficient for agreement strength at time t is the sum of its linear coefficient (-0.074) and the product of the natural log of time t and the coefficient for the interaction term (0.012). This means that, at time 0 (immediately after the war ends), the implied coefficient for agreement strength is negative, but as time passes, the implied coefficient grows larger.

[37] Greene 2003, 783.

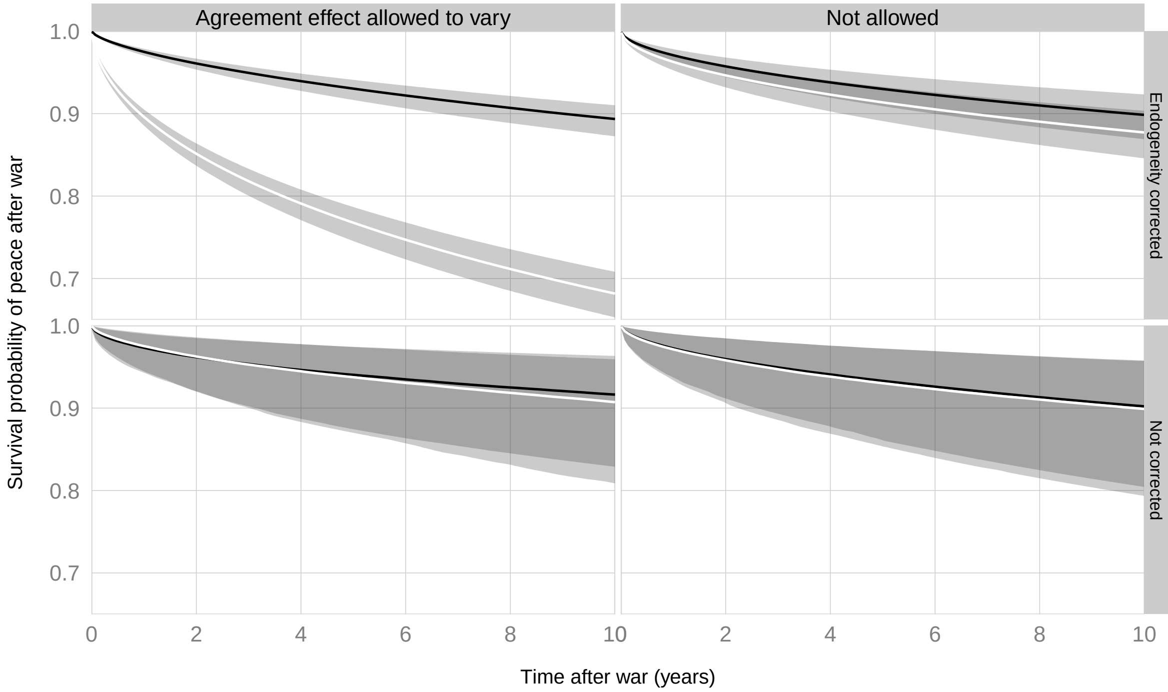

Figure 2: Substantive effects of agreement strength

Notes. This figure illustrates the estimated impact of agreement strength on postwar peace duration according to four models presented in Table 1. The horizontal axis shows time after war, and the vertical axis shows the conditional probabilities of continued peace beyond time t. Black curves show the probabilities when agreement strength is set equal to 3 and white curves show those when agreement strength is set equal to 2. The top-left panel is for Model (4), the bottom-left panel for Model (2), the top-right panel for Model (3), and the bottom-right panel for Model (1).

small change in agreement strength.

Figure 2 presents four graphs that illustrate the estimated effect of agreement strength on peace duration over time according to the four models. In each of the four panels, the horizontal axis shows time after war measured in years, and the vertical axis represents the estimated probabilities of continued peace beyond time t conditional upon peace survival until $t-1$. In each panel, the estimated probabilities of peace survival over time are shown for two scenarios: black curves correspond to the case where agreement strength is relatively strong (3 out of 10) and white curves correspond to the case where agreement strength is relatively weak (2 out of 10), holding constant all the other covariates at their median value. Gray shades around the curves show the 95 % confidence intervals of the estimate.

The top-left panel of the figure shows the results based on Model (4). We can see that a small increase in agreement strength from 2 (a value just below the median) to 3 (a value just above the median) leads to a substantial increase in the likelihood of peace survival. For example, according to the top-left panel of the figure, the estimated probability that peace lasts for more than 10 years is about 70 % (=0.7) when the disputants have a relatively weak agreement, whereas it is about 90% (=0.9) when the disputants have a relatively strong agreement. Moreover, the estimated effect is growing stronger over time, as the difference between the two curves grow larger as time passes. These findings are consistent with the theoretical expectations. On the other hand, the other three panels in Figure 2 illustrate the implied effect of agreement strength on the first three models: the bottom-right panel for Model (1), the bottom-left panel for Model (2), and the top-right panel for Model (3), all of which suggest that we would mistakenly conclude that agreement strength does not exert statistically discernible effect on peace survival when we ignore the inferential problems identified in this study.

Conclusion

This research note provides a reassessment of the relationship between the strength of cease-fire agreements and the duration of postwar peace. While past studies have provided mixed support for the theoretical argument, this study proposes an empirical strategy to test the theoretical argument. Specifically, I theorize that the effect of agreement strength grows larger as time passes since the underlying conditions that had produced a cease-fire will be likely to change over time. I also argue that we need to correct for the endogeneity of agreement strength in order to correctly estimate its effect on peace. With a newly developed statistical model, this study reports empirical findings that are commensurate with these arguments.

The proposed model not only fits the data better than conventional models do but also uncovers several important relationships that have not been documented in previous studies. First, the empirical evidence suggests that agreement strength is endogenous to *unobservable* as well as *observable* correlates of postwar peace. The negative and statistically significant estimate of the correlation parameter clearly indicates that the disputants tend to sign stronger agreements when *unobservable* factors suggest to them that they will experience shorter durations of peace. This finding complements Fortna's analysis of observable determinants of agreement strength.[38] More importantly, my finding suggests that once we correct for this endogeneity appropriately, stronger cease-fire agreements are indeed found to keep peace longer. The finding that states can sign stronger peace agreements to lower the danger of war recurrence should be very encouraging to policy makers who are concerned with conflict management. If well designed agreement can enhance the prospect for peace even under difficult circumstances, then it provides policy makers with an effective and low-cost policy measure to deal with this problem.

The finding about the increasing effect of agreement strength presents an interesting contrast with the findings from conflict mediation research. Specifically, scholars of conflict mediation have reported that the effect of mediation tends to be initially positive but short-lived.[39] That is, mediated conflicts are less likely to recur in the short run, but *more* likely to recur in the long run. Beardsley argues that this is because third-party intermediaries sometimes pressure the disputants to reach a cease-fire agreement even when the disputants have not fully resolved the bargaining problem that had caused the outbreak of conflict in the first place.[40] Such "unnatural" cease-fire agreements may maintain peace as long as the intermediaries are willing to enforce them, but third-parties often lose interests soon after immediate violence is ceased. Future research on conflict management should investigate the possibility that conflict mediation and well-designed cease-fire agreements can function in a complementary fashion to produce durable peace.

38 Fortna 2004.
39 Beardsley 2008; Gartner and Bercovitch 2006.
40 Beardsley 2008.

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Appendix 1

In this appendix, the likelihood function presented in the main text is extended to incorporate time-varying covariates. I will first sketch the standard procedure to deal with time-varying covariates in continuous-time duration analysis. I then apply the same produce to the joint statistical model presented in the main text.

In continuous-time duration analysis, whenever a covariate changes its value over time for a given subject, we duplicate the observation for the subject into two. Then, the observation with the old value of the covariate is treated as right-censored, whereas the observation with the new value of the covariate is treated as left-censored. The failure and survival probabilities for the left-censored observations are expressed as conditional probabilities, i.e., failure (survival) probability conditional upon having survived the previous period during which the covariate took a different value.

Let t_p^0 denote time at which a covariate begins to take on a new value. Then, applying the Bayes' rule, we obtain the conditional survival probability for the new observation (i.e., left-censored observation) as:

$$\begin{aligned}
\Pr(t^* > t^0 | t^* > t_p^0) &= \frac{\Pr(t^* > t^0 \cap t^* > t_p^0)}{\Pr(t^* > t_p^0)} \\
&= \frac{\Pr(t^* > t^0)}{\Pr(t^* > t_p^0)}
\end{aligned}$$

and the failure probability for the new observation as:

$$\begin{aligned}
\Pr(t^* = t | t^* > t_p^0) &= \frac{\Pr(t^* = t \cap t^* > t_p^0)}{\Pr(t^* > t_p^0)} \\
&= \frac{\Pr(t^* = t)}{\Pr(t^* > t_p^0)}.
\end{aligned}$$

The results above suggest that the likelihood contribution from the left-censored observations are obtained by simply dividing the failure and survival probabilities by the probability of survival up to the left-censoring point. Applying this logic, the joint probability of agreement strength and postwar peace duration shown in (3) can be re-written as follows:

$$\begin{aligned}
\Pr\left\{(t^* = t \cap s^* = s)|(t^* > t_p^0 \cap s_i^* = s)\right\} &= \Pr(s^* = s) \times \frac{\Pr(t^* = t|s^* = s)}{\Pr(t^* > t_p^0|s^* = s)} \\
&= \Pr(s^* = s) \times \frac{\Pr(t^* = t \cap s^* = s)}{\Pr(s_i^* = s)} \Big/ \frac{\Pr(t^* > t_p^0 \cap s^* = s)}{\Pr(s^* = s)} \\
&= \Pr(s^* = s) \times \frac{\Pr(t^* = t \cap s^* = s)}{\Pr(t^* > t_p^0 \cap s^* = s)}.
\end{aligned}$$

Similarly, the joint survival probability of postwar peace and agreement strength shown in (4) can be re-written as:

$$\begin{aligned}
\Pr\left\{(t^* > t^0 \cap s^* = s)|(t^* > t_p^0 \cap s^* = s)\right\} &= \Pr(s^* = s) \times \frac{\Pr(t^* > t^0|s^* = s)}{\Pr(t^* > t_p^0|s^* = s)} \\
&= \Pr(s^* = s) \times \frac{\Pr(t^* > t^0 \cap s^* = s)}{\Pr(s^* = s)} \Big/ \frac{\Pr(t^* > t_p^0 \cap s^* = s)}{\Pr(s^* = s)} \\
&= \Pr(s^* = s) \times \frac{\Pr(t^* > t^0 \cap s^* = s)}{\Pr(t^* > t_p^0 \cap s^* = s)}.
\end{aligned}$$

Appendix 2

Table A.1: Agreement strength: Constituent variables and their measurements

Variable Name	Measurement
Formalism	0 = no declared cease-fire, or tacit or informal acceptance of cease-fire 1 = formal acceptance of cease-fire proposal or agreement
Withdrawal of Forces	0 = none 1 = partial, to status quo ante, or beyond
Demilitarized Zones	0 = none 1 = demilitarized zone
Arms Control	0 = none 1 = arms embargo, limits near cease-fire line, specific weapons prohibited
Peacekeeping	0 = none .5 = monitoring (unarmed military observers) 1 = peacekeeping forces (armed)
External Involvement	0 = none .5 = mediate cease-fire, restraint, patron, etc 1 = explicit or well-understood guarantee of peace
Paragraph Count	0 = 0 paragraph 1/3 = 1–20 paragraphs 2/3 = 21–80 paragraphs 1 = over 80 paragraphs
Internal Control	0 = none .5 = stated responsibility for actions from own territory 1 = concrete measures to ensure control
Confidence Building Measures	0 = none 1 = military info exchanged, hot line, onsite or aerial verification
Dipute Resolution	0 = none .5 = ongoing third-party mediation 1 = joint commission of belliegerents

Source: Fortna (2003, 2004).
Index of agreement strength is constructed by summing all these ten variables. In this study, I multiple the index by 6 so that it only takes non-negative integer values while keeping the substantive interpretation of individual contributions of constitutive variables intact.