journal of peace RESEARCH

© 2006 Journal of Peace Research,
vol. 43, no. 3, 2006, pp. 243–260
Sage Publications (London, Thousand Oaks, CA
and New Delhi) http://jpr.sagepub.com
DOI 10.1177/0022343306063930

Interdependence and the Duration of Militarized Conflict*

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Many researchers have reported empirical support for the liberal proposition that increased trade between states reduces their propensity to engage in militarized conflict. However, the literature has been less vocal on the effects of interdependence once actual conflict has started. The author, who builds on the opportunity-cost explanation for the commercial peace, first shows how that explanation can fit a bargaining framework of conflict. It is then argued that if trade reduces conflict by raising its opportunity costs prior to its onset, the prospective opportunity costs should remain high during conflict as well and decrease its observed duration, a prediction that can also be derived from a simple war-of-attrition game. The proposed inverse relationship between interdependence and conflict duration is evaluated empirically, using all militarized interstate disputes between 1950 and 1992. The author utilizes both continuous and discrete-time duration models and considers a number of alternative model specifications and concept measurements, and the results are generally supportive of the tested hypothesis. In addition, a two-stage model of dispute onset and duration is estimated, which reveals that more probable disputes are also longer, but the observed selection pattern does not seem to influence the positive effect of interdependence on the hazards of dispute termination.

Introduction

Many researchers have been attracted by the classical liberal proposition that extensive trade prevents militarized interstate conflict. Indeed, the hypothesis that economic interdependence has a pacifying effect on international relations has been successfully tested at the dyadic level (e.g. Oneal & Russett, 1997, 1999; Russett & Oneal, 2001). However, while scholars have frequently looked into whether higher dyadic trade decreases the odds of a militarized interstate dispute (MID) onset, less attention has been devoted to the potential effects of trade on conflict that has already begun.[1] I, on the contrary, argue that if economic interdependence inhibits conflict, in particular by raising its opportunity costs, these opportunity costs should remain high during conflict as well and contribute to its faster resolution. I deduce the latter proposition from the opportunity-cost argument, or one of the two major explanations for the commercial peace, and subject it to a series of empirical tests that indeed indicate an inverse relationship between interdependence and the duration of militarized conflict.

* The author wishes to thank Erik Gartzke, T. Clifton Morgan, Brett Ashley Leeds, Randolph Stevenson, Richard Stoll, William Reed, Mark Crescenzi, and three anonymous reviewers for their helpful comments. The replication dataset is available at http://www.prio.no/jpr/datasets. Correspondence should be directed to vulcho@rice.edu.

[1] An exception is Levy (2003), who addresses in brief the possibility that trade also provides incentives for conflict termination. Gartzke (2003a), on the other hand, discusses the effect of trade on conflict escalation.

Although previous studies have not inquired explicitly into the effect of trade on conflict duration, a voluminous literature can be located under the more general heading of whether trade reduces conflict, and I present a brief review of this work below. I build on the opportunity-cost argument, and, in view of some recent criticisms, I focus on its crucial assumption that war has an adverse effect on commerce. Following that, I use a simple war-of-attrition model to formalize the intuition for why interdependence, conceptualized as the higher opportunity costs within a dyad, should decrease the expected duration of militarized conflict. I then evaluate the latter proposition through a series of duration models applied to all MIDs between 1950 and 1992. I consider alternative specifications and estimate a two-stage model to control for the potentially non-random selection when dyads cross the MID threshold. Generally, the results support the theoretical expectation. I conclude by discussing the implications of my findings for the broader study of trade and conflict.

The Study of Trade and Conflict

Multiple causal mechanisms explaining how economic interdependence inhibits international conflict have been put forward in the literature. Examples include the arguments that trade and military conquest are alternate means of acquiring scarce resources, more efficient trade thus making conflict less attractive, and that economic cooperation between private actors eventually spills over into the political domain (Mansfield & Pollins, 2001).[2] A frequently recurring claim has been that commerce enriches domestic and foreign private agents, whose welfare constrains government action, as conflict might disrupt trade (Polachek, 1980; Mansfield & Pollins, 2001).[3]

According to Polachek, Robst & Chang (1999: 405), the simple logic behind the pacifying effect of trade is that if conflict disturbs trade, 'then countries with the greatest gains from trade face the highest costs of potentially lost trade and hence engage in the least conflict and most cooperation'. However, while the cited argument exemplifies the 'opportunity cost' hypothesis for the link between trade and conflict (Mansfield & Pollins, 2001; Gartzke, 2003b), there are also signaling arguments that view trade as an informational medium allowing states to signal their resolve (e.g. Gartzke, Li & Boehmer, 2001; Gartzke, 2003a,b).

Both opportunity-cost and signaling arguments lead to the expectation that trade and conflict would be inversely related, and, indeed, the empirical literature has been quite supportive of that claim (Polachek, 1980; Gasiorowski & Polachek, 1982; Oneal et al., 1996; Oneal & Russett, 1997, 1999; Russett & Oneal, 2001). A number of scholars have also provided qualified support for the statistical relationship between trade and conflict. For instance, Gasiorowski (1986) has argued that trade interdependence has both conflict-promoting and cooperation-promoting aspects. Mansfield & Pevehouse (2000), in turn, report that the pacifying effect of interdependence is contingent upon states' membership in the same preferential trading agreement, while Hegre (2000) argues that the effect of interdependence is interactive with the level of economic development. Finally, Crescenzi

[2] Hegre (2004) presents a formal model built on the substitutability argument. The spillover mechanism has been extensively explored in the neo-functionalist literature.

[3] As Simmons (2003) points out, that view presupposes some domestic pluralism that allows discontented private agents to exert pressure on their government's foreign policy. However, autocratic governments might directly benefit from trade, as the cases of some oil-exporting countries suggest.

(2003) shows that trade decreases high-level but increases low-level conflict.[4]

These positive empirical findings have been challenged on several fronts. First, Barbieri (1996, 2002) has questioned their validity by showing that interdependence increases the likelihood of a MID onset in a given dyadic year. Oneal & Russett (1997), nevertheless, have raised as an objection the fact that Barbieri's studies include all possible dyads, an arrangement which might produce a spurious correlation between trade and conflict, as contiguous states both trade more and fight more often. In addition, Gartzke & Li (2003) have shown that Barbieri's measure of trade share is negatively correlated with trade openness, which captures the importance of international trade to a state's economy. Beck, Katz & Tucker (1998), in turn, have argued that Oneal & Russett's (1997) strong results in support of the commercial peace arise only because of their failure to control for temporal dependence. Still, Oneal & Russett (1999) change their model specification in accordance with Beck, Katz & Tucker's (1998) recommendations and again find statistical evidence for the pacifying effects of trade.

A strong criticism of the commercial peace arises from the observation that firms, not governments, are responsible for most commercial activities (Morrow, 1999). If firms with rational expectations anticipate conflict, they will try to avoid the business risks associated with war by divestment of their endangered foreign assets or looking for alternative trading partners. Thus, trade levels should fall before conflict and have little impact on its onset. Barbieri & Levy (1999), on the other hand, find that war does not affect trade as seriously as many scholars have assumed. Even if war decreases dyadic trade in the short run, it usually has no long-term adverse effect, and, in many instances, the volume of postwar trade exceeds the prewar volume. Moreover, there have also been historical cases in which bilateral trade has continued even during wartime (Barbieri & Levy, 1999).[5]

Both Morrow (1999) and Barbieri & Levy (1999) undermine the commercial peace literature. However, while Barbieri & Levy doubt the empirical validity of the critical opportunity-cost assumption that conflict has an adverse effect on trade, Morrow's argument implies the theoretical inconsistency of that assumption. A potential reconciliation with the positive empirical findings in support of the commercial peace is proposed by Li & Sacko (2002), who present a more comprehensive theory of conflict expectations than Morrow's. Li & Sacko (2002) emphasize that Morrow's argument relies on the controversial premise that firms have perfect foresight with respect to future conflict. However, if war indeed arises because of asymmetric information, and its onset is always probabilistic, as recent theories claim (e.g. Fearon, 1995; Gartzke, 1999), then assuming firms have perfect foresight would imply they know more about future conflict than do their respective governments. Moreover, firms' lack of perfect foresight is also suggested empirically by Barbieri & Levy (1999: 474–475), who

[4] Crescenzi (2003) also argues that the concept of interdependence should ideally capture the difficulty in finding an alternative trade partner, rather than the amount of trade. A state's imports of some strategic material (e.g. uranium) might be negligible compared to its total trade, but finding an alternative supplier might be impossible, whereas the large-scale imports of consumer goods might be easily replaceable (Ripsman & Blanchard, 1996/97; Mansfield & Pollins, 2001; Crescenzi, 2003). Crescenzi measures interdependence by interacting trade volume with price elasticity, which measures a state's sensitivity to import price fluctuations, but only covers a limited spatial and temporal domain owing to data limitations. I follow Oneal & Russett and use trade volumes to measure interdependence, implicitly assuming they are positively related to the utility from trade.

[5] Anderton & Carter (2001), however, look at different warring dyads from Barbieri & Levy (1999) and find that war reduces trade significantly, their findings being stronger for major than for minor power dyads.

acknowledge, concerning the prewar periods in their study, that 'there is no systematic evidence that the anticipation of war usually results in a reduction of trade'.[6]

Li & Sacko (2002) relax the perfect foresight assumption and introduce the expectedness of conflict as a variable influencing pre-conflict trade. They agree that firms would find it preferable to relocate their business prior to conflict when its onset is expected, but argue that when conflict is unexpected, firms are less likely to be prepared in advance, and most of the business relocation should occur after its onset, leading to the empirical observation that conflict reduces trade. Li & Sacko also offer empirical support for their argument, as they show that unexpected MIDs reduce trade volumes more than expected MIDs do.

Firms' imperfect foresight allows militarized conflict to reduce trade, as the validity of the opportunity-cost argument requires. However, even perfect firm foresight might not be able to eliminate the economic opportunity costs of conflict. First, if firms can anticipate their governments' foreign policy, governments should also be capable of correctly predicting the firms' behavior. If governments, in turn, are aware of firms' intentions to divest when the probability of conflict becomes high, they will take the expected pre-conflict trade reduction into account even before adopting a policy course that would make the firms see conflict as likely in the first place. Higher trade volumes would again make conflict unattractive, the only difference from the common form of the opportunity-cost argument being that governments' cost–benefit calculations involving trade would be simply pushed back to an earlier stage of the given crisis.

Second, even if conflict starts after firms have already relocated their business, the disputing states should still face some prospective opportunity costs. Provided that firms trade with their most preferred partners, the relocation of business activities to second-best trading partners should still fall short of total compensation for the lost exchange. Even if the rigidity of firm practices can be eventually overcome, in the short term firms will be less efficient if they cannot trade with their original partner, regardless of whether that trade has been interrupted before, during, or after conflict onset.[7]

Finally, while the argument addressed here explicitly concerns the pre-conflict divestment of firms, trade is not always conducted by firms. In centrally planned economies, it is the state that is responsible for all trade, and, even in market-oriented economies, states sometimes command a significant public sector that trades internationally. The opportunity-cost logic can be undermined if, owing to firms' divestment prior to conflict onset, by the time governments have to decide whether to fight, trade is already lost and plays no role in their calculations. However, when a state controls its trade directly, that trade will remain in consideration because, while the state might have a hard time convincing its firms to reinvest if it decides to avert conflict, it can easily restore its own commercial operations. That immediate possibility of restoring trade, in turn, will always guarantee the presence of prospective opportunity costs due to interdependence. Thus, it is just as unclear whether firms' rational expectations contradict the opportunity cost logic as it is unclear whether firms can have such expectations in the first place. That is why I assume that trade raises the opportunity costs of conflict, and, in the next section, I show how the

[6] Anderton & Carter's study questions Barbieri & Levy's conclusions about trade declining during, rather than before, war.

[7] The prospective (that is, still not incurred) losses due to the decreased firm efficiency should translate into higher opportunity costs of conflict for the disputing states at any point of time. Therefore, trade should reduce conflict both before and after its onset, as I argue in the next section.

inverse relationship between trade and the duration of conflict can be deduced from that assumption.

Interdependence and the Duration of Conflict

According to the opportunity-cost argument, interdependence promotes peace by raising the costs of militarized conflict (Polachek, 1980; Polachek, Robst & Chang, 1999). Conflict becomes more costly, in turn, because the fighting parties, in addition to bearing the costs of waging warfare, forfeit the potential gains from trading, owing to government-imposed restrictions and increased business risks. However, these conflict-inhibiting effects of interdependence are not limited only to the pre-conflict phase of a dispute, and the opportunity-cost argument can explain how the prospect of further trade losses provides incentives for conflict termination as well.

As some scholars have observed, any theory of the effect of interdependence on conflict should be grounded in a solid understanding of the occurrence and dynamic of conflict itself (Morrow, 1999, 2003; Gartzke, 2003b). While traditionally multiple theories of conflict have proliferated in the study of IR, recent scholarship has drawn attention to its informational origins (Fearon, 1995; Gartzke, 1999). As Fearon (1995) argues, if most conflicts end in some negotiated settlement over the disputed issue, rational states should prefer to conclude that settlement prior to incurring the conflict costs, as the bargaining range of mutually acceptable settlements is guaranteed to be non-empty when these costs are positive. A very common reason for states sometimes being unable to reach a rational pre-conflict settlement emerges in the asymmetry of information, combined with states' incentives to misrepresent their reservation values. Conflict, on the other hand, helps states to credibly communicate these reservation values by demonstrating their willingness to incur its costs or revealing the true magnitude of the costs, as an expanding informational literature on war suggests (e.g. Wagner, 2000; Filson & Werner, 2002; Slantchev, 2003).

The opportunity-cost logic implies that interdependence can enter the theoretical framework outlined above through the conflict-cost parameters, as interdependence increases these costs. Following Fearon's (1995) discussion, higher conflict costs increase the pre-conflict bargaining range and should, therefore, decrease the probability of conflict. In their calculus, states balance the size of their demands against the probability that these demands exceed the opponent's reservation value and are rejected. Higher conflict costs due to greater interdependence worsen states' conflict payoffs and push them to lower their demands, which, in turn, results in a reduced probability of conflict onset.[8]

Signaling arguments, on the other hand, suggest that interdependence allows states to credibly communicate their resolve or reservation values by severing an advantageous economic relationship that an unresolved state would not terminate. The credible communication made possible by interdependence reduces the uncertainty existing over the bargaining range and increases the likelihood of a settlement short of war (e.g. Gartzke, 2003a,b; Morrow, 2003). Thus, if

[8] The inverse relationship between the costs and probability of conflict can also be derived formally from strategic interaction models, although the result is not insensitive to the model specification. Reed (2003) analyzes an ultimatum model with one-sided uncertainty over the distribution of power, in which an increase in any state's conflict costs leads to a reduced probability of disagreement. Coletta & Gartzke (2003) analyze an ultimatum model with one-sided uncertainty over the costs of the demand respondent and find that, while the probability of conflict is always decreasing in the costs of the demand sender, the effect of the respondent's costs is contingent upon the assumptions one makes about the belief distribution function.

we adopt Fearon's (1997) terminology, signaling implies that interdependence allows states to 'sink costs', while the opportunity-cost logic is more reminiscent of 'tying hands'; that is, interdependence affects states' behavior by changing their incentives.

The opportunity-cost argument for why interdependence inhibits militarized conflict can be easily extended to account for the effect of interdependence on the duration of conflict. If interdependence raises the opportunity costs of conflict prior to its onset, then these costs should also remain high after onset, because, at least in the short term when firms have not permanently reoriented their business operations, they will gain if hostilities cease and normal trade with the adversary is restored. Then, just as the higher prospective costs of conflict push states to lower their demands and avert conflict prior to its onset, so do these higher prospective costs push states to settle early, even if conflict has not fully served its informational purpose and states might be forfeiting the better deal they can get if they know more. That is, the purpose of militarized conflict is to overcome asymmetric information, but conflict costs are the price states have to pay to extract that information. The higher these costs are due to interdependence, the more expensive the information-revelation process is, and the sooner are states likely to settle on unfavorable terms rather than continue fighting.

To demonstrate formally how higher opportunity costs for either disputant in a militarized conflict can reduce the expected duration of that conflict, I use a war-of-attrition model (Fudenberg & Tirole, 1991: 119–126). In a war-of-attrition, two players compete for an indivisible prize while incurring costs at a constant rate. At any time, each player has two available actions, namely to exit the competition or to hold on. When a player exits, the game ends and the opponent wins the prize, but both players incur the competition costs. Although the war-of-attrition presented here is one of complete information and the issue at stake is indivisible, conflict arises in equilibrium, owing to the use of mixed strategies, which is why this simple game can serve as an approximation for a more flexible model that allows for endogenous demands.[9] Again, the purpose of the war-of-attrition analyzed below is not to generate novel insights, but rather to capture the intuition that the costs and duration of conflict should be inversely related.

In my example, assume that two players, A and B, are contesting some issue worth v to both of them, with the status quo set to (0,0). Let A be the revisionist player, so that if A wins the contest, the new allocation will be $(v,-v)$, while if A loses, the status quo will remain. The time t in the game is continuous and ranges from 0 to $+\infty$. A and B incur costs at the rate of $c > 0$ per unit of time, and at each time $t \in [0, +\infty]$ they simultaneously choose an action $a \in \{stay, exit\}$, given the other player has not exited before. As soon as someone exits, the game ends and the player to exit first forfeits its preferred allocation.[10]

As Fudenberg & Tirole show, such games have a unique symmetric equilibrium, which is in mixed strategies that make each player indifferent between exiting and waiting for an infinitesimal amount of time at any point of the game. Since the players' payoffs are symmetric, let their equilibrium strategy be represented by the cumulative density function $F(t)$, which gives the probability that a player has exited by time t. A's strategy has to make B indifferent between exiting and staying at any time for B to be able to

[9] Furthermore, the described mixed-strategy equilibrium is also the limit of a series of incomplete information war-of-attrition equilibria, in which each player type is associated with a distinct pure strategy (Fudenberg & Tirole, 1991: 230–232).

[10] Since the described equilibrium is in mixed strategies and time is continuous, it is inconsequential what happens if the two players exit simultaneously.

play a mixed strategy, and, in turn, *A* must be indifferent too, owing to *B*'s strategy. For each player to be indifferent between exiting and staying a little longer, the marginal benefit of staying should equal the marginal cost. That is, the value of the stakes for each player multiplied by the probability that the other player exits in some arbitrarily small time period *dt* should equal the player's costs incurred during *dt*. Therefore, the equilibrium $F(t)$ is the solution to the following equation:

$$v\ \frac{dF(t)/dt}{1-F(t)} - dc(t)/dt = 0. \qquad (1)$$

The first term in the equation is the hazard rate of the opponent's mixed strategy distribution, or the instantaneous probability that the opponent will exit given that he has not exited until time *t*, multiplied by *v*, which is the difference between winning and losing for both players. The second term is simply the marginal cost for the respective player. Since the linear costs imply a constant hazard rate, Equation (1) is solved by the exponential distribution, that is, $F(t) = 1 - e^{(-ct/v)}$. This exponential distribution gives the probability that a player has exited the game by time *t*.[11]

However, since we are ultimately interested in the expected duration of conflict, we need the distribution governing the event that any player exits by time *t*. If we denote that distribution by $H(t)$, for any *t*

$$H(t) = 2F(t) - F(t)F(t) = 1 - e^{(-2ct/v)}. \quad (2)$$

Since $H(t)$ is also exponential, the expected duration is equal to the inverse hazard rate, or

$$E(t) = \frac{v}{2c}. \qquad (3)$$

[11] The described equilibrium is also subgame perfect, because the play is always on the equilibrium path when the game has not ended.

It is easy to see that any increase in the marginal costs leads to a decrease in the expected duration of conflict, and we have already conceptualized interdependence as a relationship in which both sides have higher opportunity costs. Therefore, adding the costs associated with interdependence to the costs of militarized conflict should decrease its expected duration.[12]

While this extension of the opportunity-cost logic suggests that interdependence reduces the duration of conflict, the inferences from signaling arguments are more ambiguous. Morrow (2003) argues that costly signals of resolve, for example the imposition of trade sanctions, should be used mostly prior to crisis militarization and less frequently after MID onsets. Indeed, if a dispute still occurs after signaling, it must be that the signaled resolve has been insufficient, and further trade signals are unlikely to succeed, because they should also fall short of signaling the resolve necessary to avert conflict. Therefore, if interdependence pacifies through richer signaling menus, after dispute onsets it should have an indeterminate role on dispute duration, as getting involved in a dispute already signals a greater resolve than interdependence can possibly convey.[13]

Levy (2003), on the other hand, suggests that, if the opportunity-cost logic holds, interdependent dyads that have ended up in a dispute should have been selected on their

[12] This inverse relationship is also demonstrated by Kennan & Wilson (1989: S100), who simulate the outcomes of a more complicated, incomplete information, war-of-attrition in the context of labor strike duration.

[13] Gartzke (personal communication with the author) suggests that viewing conflict as a repetition of the one-shot escalation games analyzed by Gartzke (2003a) leads both signaling and opportunity-cost arguments to predict shorter conflict when interdependence is greater. He argues that, since signaling implies more credible communication, interdependent disputants should require fewer repetitions of the one-shot game to eliminate the asymmetric information between them. Opportunity costs, on the other hand, imply selection on resolve and higher attrition rates for interdependent disputants.

greater resolve, which should increase dispute duration and intensity. Signaling predicts shorter dispute duration when interdependence is higher, according to Levy (2003), owing to the existence of more credible information about the disputants' reservation values. Nevertheless, while Levy's claim that dyads that have crossed the dispute threshold are a non-random sample selected on greater resolve is quite plausible, the relationship he proposes between interdependence and conflict duration does not follow automatically from this selection effect.

It is very likely that issue salience for dyads involved in MIDs is correlated with conflict onset, since rational states will start a conflict only if their expected gain exceeds the expected costs. However, the higher issue salience or unobserved resolve for those dyads does not imply that, conditional on a dispute occurring, the interdependence costs will prolong conflict. Selection on resolve, information, or other unobservables can certainly confound the relationship between conflict costs and duration. Yet, determining the magnitude of that selection effect is an empirical problem, and there is no theoretical reason for higher costs having opposite effects before and after conflict onset. Nonetheless, Levy's (2003) argument emphasizes the importance of carefully controlling for the non-random selection of dyads into militarized disputes when an inference is sought about an observable aspect of these disputes, and I do address that issue in greater depth in the empirical section.

Thus, if one accepts the opportunity-cost premise that trade adds additional costs to conflict in terms of foregone revenue, it should also be the case that more interdependent states face greater opportunity costs during conflict as well. These greater costs, in turn, should provide incentives for more expedient conflict termination as they raise the cost states have to pay to fight, a hypothesis that I subject to a series of empirical tests below.[14]

Research Design

The trade-and-conflict literature has frequently operationalized military conflict as a MID to the advantage of significantly more conflict cases compared to full-scale wars. I also identify military conflict with the occurrence of a MID, defined as a historical case of conflict in which one state directed the threat, display, or use of military force against another state (Jones, Bremer & Singer, 1996). Owing to the limited availability of trade and GDP data before 1950, my population of cases includes all MIDs that started between 1950 and 1992 drawn from Maoz's Dyadic MID Dataset (Maoz, 1999).[15] I measure the dependent variable of MID duration using the number of days elapsed between the dyadic MID start and end dates as adjusted by Maoz.[16]

Although I consider Barbieri's (1996) measure, in most model specifications below interdependence is defined as the lower

[14] Levy (2003) points to the war of the Spanish Armada as an example of how forgone trade pressures state leaders to terminate an ongoing conflict. He draws from Croft (1989: 301), according to whom 'Merchants and mariners who made a living from the Iberian trades viewed the drift to war with gloom and wanted a speedy end to hostilities.' By the end of the 16th century, both Elizabeth and Philip II were facing severe economic difficulties, which eventually forced them to cease the hostilities, the prospect of unrestricted trade being one of the strongest arguments for ending the war (Croft, 1989: 297).

[15] I use the updated version of the dataset available as 'dyadmid602.csv' from EUGene (Bennett & Stam, 2000).

[16] Calculating MID duration is straightforward for bilateral MIDs with exactly one disputant on each side. However, the start and end dates of multilateral MIDs do not always reflect the duration of the component dyadic MIDs, owing to late entries, early exits, or the lack of direct conflict. The problem has been solved by Maoz (1999), who uses the original incident-level data to eliminate invalid dispute dyads after disaggregating the multilateral MIDs. The 16 dyadic MIDs coded by Maoz as ending on the last covered date, 31 December 1992, have been treated as censored spells except for MID #3554, which indeed ends then.

dyadic trade-to-GDP ratio in each dyad (Oneal & Russett, 1997), and I have drawn both the trade and GDP data from version 4.1 of Gleditsch's Expanded Trade and GDP dataset (Gleditsch, 2002). Several variables might confound the relationship between interdependence and MID duration. First, it is important to control for distance, because states that are close geographically experience both more conflict and more cooperation (e.g. Oneal & Russett, 1997).[17] To control for geographical proximity, I use *ln(distance)*, defined as the natural logarithm of the distance between the disputants' capitals, and *contiguity*, which is a binary variable coded 1 if the disputants were land contiguous or separated by less than 150 miles of water and 0 otherwise. *Ceteris paribus*, both smaller distances and contiguity should prolong dispute duration.[18]

Second, I control for the power disparity between the disputing states. Bennett & Stam (1996), who analyze war duration, argue that power disparity decreases war duration because stronger states need less time to inflict sufficient damage to force a weaker state to surrender. On the other hand, Bueno de Mesquita, Koch & Siverson (2004), who analyze MID duration, find that power disparity is a good predictor of dispute duration only for democratic dyads. Democratic leaders choose to participate in disputes which they are likely to win, while autocratic leaders are less concerned about losing a dispute and less likely to take power disparity into account. Bueno de Mesquita, Koch & Siverson (2004), however, look only at prewar dispute duration and truncate MID spells after the war threshold is crossed, whereas I treat both war and minor disputes as forms of militarized conflict and use the complete MID spells. *Ln(power disparity)* is set to the natural logarithm of the ratio of the higher over the lower Correlates of War capability score in the disputing dyad (Singer, Bremer & Stuckey, 1972) and should, *ceteris paribus*, decrease MID duration.

I also control for a disputing dyad's alliance status and joint democracy. Alliances are proxies for common interests, generally found to inhibit MID onsets (e.g. Oneal & Russett, 1999). Moreover, Bueno de Mesquita, Koch & Siverson (2004) show that alliances decrease MID duration, which fits the war-of-attrition logic if fighting an ally is more costly, owing to negative externalities. *Alliance* is coded 1 if a dyad is formally allied and 0 otherwise (Gibler & Sarkees, 2002) and should decrease MID duration. Joint democracy might also reflect the higher costs of fighting for democratic dyads (e.g. Bennett & Stam, 1996), but the existing theoretical explanations for how democracy inhibits conflict are simply too many to be discussed here. It should suffice that Bueno de Mesquita, Koch & Siverson (2004) demonstrate empirically that democratic disputes are indeed shorter, which suggests that democracy is an important covariate of MID duration. *Joint democracy* is coded 1 if both disputants have a Polity IV democracy minus autocracy score of 6 or greater (Marshall & Jaggers, 2002) and 0 otherwise.[19] Jointly democratic dyads should experience shorter MIDs.

I first estimate the effects of independent variables on MID duration using the semi-parametric Cox proportional hazards model, which is more flexible than parametric models as it does not make assumptions about the shape of the baseline hazard rate.[20]

[17] Also, conducting military operations at a distance is more difficult and more costly (e.g. Bremer, 1992), which is why greater distance should be a proxy for higher conflict costs and decrease MID duration following the war-of-attrition logic.

[18] Capitol-to-capitol distance, contiguity, and the other control variables are drawn from EUGene.

[19] I use the Polity 2 score.

[20] I follow the discussion by Box-Steffensmeier & Jones (2004). The war-of-attrition implies flat hazards, but I view it as a formalization of an intuitive claim rather than as an exact model of MID dynamics.

Following that, I complement my analysis through the discrete-time Prentice–Gloeckler–Meyer (PGM) model (Jenkins, 2004–05). Finally, I estimate a two-stage model of MID onset and duration, in which the predicted probability of a dyadic MID from the first stage serves as a covariate of MID duration at the second stage. In total, I report the results from nine different specifications, all discussed below.[21]

Data Analysis

Since the parameterization in both the Cox and discrete-time models I use is in terms of the hazard rate, a positive coefficient implies that the hazard rate or risk of termination increases for higher values of a covariate. Higher hazard rates imply shorter duration, so a positive coefficient means that a covariate decreases the time until failure. Conversely, a negative coefficient implies a decrease in the hazard and an increase in the time until failure. The exponentiation of a coefficient gives the ratio of the hazard rate when the corresponding covariate has a value of 1 over the baseline hazard when all covariates are set to 0 (Box-Steffensmeier & Jones, 2004). Table I reports the Cox estimates of MID duration for five specifications and the discrete-time estimates for three additional specifications.[22]

Model 1 is my initial specification, in which I regress MID duration on the lower trade-to-GDP ratio and the five main controls: *ln(distance)*, *contiguity*, *ln(power disparity)*, *alliance*, and *joint democracy*. As Table I shows, the coefficients of all covariates except for contiguity have the anticipated signs and are statistically significant. More specifically, the null hypothesis that interdependence, measured as the lower dyadic trade-to-GDP ratio, does not affect MID duration can be rejected at the $p = 0.004$ level, which supports the argument presented in the theoretical section.[23] As for contiguity, Model 1 shows surprisingly that it is associated with shorter MIDs. Although I do not have a theoretical explanation for this result, it is likely that most of contiguity's effect is appropriated by the log of distance, owing to the high correlation between the two variables.[24]

The Cox model assumes that the effect of the covariates on the hazard ratio is proportional over time; a dyad's alliance status, for instance, should have a constant impact throughout a MID spell. Since the proportional hazards assumption is frequently violated in IR research, it is important to check its validity and correct for non-proportional effects by interacting the 'offending covariates' with an appropriate transformation of time (Box-Steffensmeier & Zorn, 2001; Box-Steffensmeier, Reiter & Zorn, 2003). Following a procedure suggested by the aforementioned authors, I conduct a proportionality test, which reveals the non-proportional effects of *alliance* and *joint democracy*.[25]

Model 2 amends for this non-proportionality by including interactions of the two offending covariates with the natural log of time. The time-fixed effect of *alliance* now appears insignificant while its positive continuous-time-varying coefficient suggests that as MIDs continue, alliances exert stronger pressure on their termination. *Joint democracy* retains its positive coefficient, but

[21] All statistical analysis is conducted in Stata 8.2 (StataCorp, 2003).

[22] I use the Efron method for dealing with ties in the Cox models. These were obtained in Stata through *stcox* with the *efron* option. All models use time-fixed covariates measured in the year preceding the MID. I also estimate a model with time-varying covariates, which is reported in the accompanying log file, but is omitted here as the results are essentially the same.

[23] Estimating a control model without the lower trade-to-GDP ratio results in practically the same coefficients for the control variables.

[24] *Contiguity* also appears insignificant when only originator dyads are considered.

[25] I use the Grambsch & Therneau test or *stphtest*, *detail*, replacing time by its natural log as recommended by Box-Steffensmeier & Zorn (2001).

has a negative interaction with time, which suggests that, early in MIDs, democracy facilitates MID termination, but later on its influence weakens.[26] Most importantly, however, the effect of interdependence is proportional and retains its significance in the improved specification. Increasing the lower trade-to-GDP ratio by one standard deviation from its mean increases the baseline hazard by 6%, while increasing it from its minimum of 0 to its maximum in-sample value of 0.147 increases the hazard by 315%. This effect compares well to the maximum effects of *ln(distance)* (+74%), *contiguity* (+19%), *ln(power disparity)* (+44%), and *joint democracy* (+152%).

Following Box-Steffensmeier & Jones (2004: 124–130), I perform additional diagnostics on Model 2. Plotting the Cox–Snell residuals against their integrated Kaplan–Meier hazard rate indicates a good overall fit, as the plot follows the 45-degree line through the origin. To check for influential observations, I also estimate the model multiple times, dropping each observation once, but do not find evidence that a particular observation is driving the results.[27] To check for non-linear effects, I plot the smoothed martingale residuals from Model 2 against each covariate (see Box-Steffensmeier & Jones, 2004: 127). I do not find clear aberrations, except maybe for the effect of interdependence itself, which is slightly better captured by a square root transformation.[28]

According to Box-Steffensmeier, Reiter & Zorn (2003), an important issue is that of unobserved heterogeneity, especially when duration models involve repeated events in a sub-sample category. The aforementioned authors' advice is relevant to MID duration, as many dyads have experienced multiple MIDs between 1950 and 1992. It might be, for instance, that dyads characterized by greater overall hostility experience longer disputes. To control for unobserved heterogeneity, I also estimate the Cox regression reported as Model 3 with a frailty parameter, or a random effect that attempts to characterize the unobserved failure propensity of MID spells, shared at the dyadic level (Box-Steffensmeier & Jones, 2004: 142–148).[29]

Model 4 convincingly shows the presence of unobserved heterogeneity, and most covariates lose their statistical significance. The lower trade-to-GDP ratio retains marginal significance, which suggests that the variable is not completely insensitive to unobserved heterogeneity. I was unable, however, to isolate the source of that unobserved heterogeneity, as plotting the disputing dyads against their associated frailties did not suggest any concrete omitted variable as that source. In fact, the heterogeneity might be due to inherently immeasurable variables such as resolve, which significantly complicates the problem. Still, the correct sign and marginal significance of the interdependence coefficient remain supportive of the theoretical prediction.

In Model 4, I use Barbieri's (1996) interdependence indicator. The corresponding coefficient is significant, which shows that the result is robust to the main alternative of the lower trade-to-GDP operationalization.[30] Model 5 controls for a dyad's lower trade openness, or total trade-to-GDP ratio, and again supports the central result. In

[26] Stata's *tvc* option was used to include the two variables' interactions with time in Model 2. The reported non-proportionality was also assessed through a likelihood ratio test between Model 2 and the nested Model 1.

[27] I perform the procedure on Model 1 since the *tvc* option in Model 2 makes it too computationally intensive. The only clearly influential observation decreases the interdependence coefficient; that is, omitting it increases the support for the tested hypothesis.

[28] The accompanying log file reports a Cox model that uses that transformation. Although there is some improvement in the log-likelihood, I report the linear effect to simplify the interpretation of the results.

[29] I add the *effects(.)* and *shared(.)* options to the *stcox* command in Stata.

[30] Barbieri's measure is defined for the AB dyad as Sqrt((Trade AB/Total trade A)* (Trade AB/Total trade B))*(1–|(Trade AB/Total trade A)– (Trade AB/Total trade B)|). The results are similar when each country's total trade is substituted with its GDP.

Table I. The Effect of Interdependence on the Hazards of MID Termination, 1950–92[a]

Independent variable	*Model 1*	*Model 2*	*Model 3*	*Model 4*	*Model 5*	*Model 6*[b]	*Model 7*	*Model 8*[c]
	Cox	*Cox*	*Cox*	*Cox*	*Cox*	*PGM*	*PGM*	*PGM*
Ln(distance)	0.061**	0.059**	–0.022	0.056**	0.077***	0.055**	0.197	0.213
	(0.027)	(0.027)	(0.042)	(0.026)	(0.028)	(0.027)	(0.244)	(0.243)
Contiguity	0.178**	0.177**	0.111	0.201***	0.214***	0.185**	0.111	0.082
	(0.073)	(0.073)	(0.098)	(0.073)	(0.074)	(0.074)	(0.629)	(0.627)
Ln(power disparity)	0.042**	0.039**	0.043*	0.036*	0.045**	0.064***	–0.062	–0.033
	(0.019)	(0.019)	(0.025)	(0.019)	(0.019)	(0.019)	(0.178)	(0.177)
Alliance[d]	0.219***	–0.045	–0.119	–0.068	–0.069	1.481***	–0.379	–0.491
	(0.068)	(0.110)	(0.123)	(0.111)	(0.111)	(0.144)	(0.626)	(0.625)
Joint democracy	0.607***	0.922***	0.919***	0.896***	0.928***	2.400***	6.879***	6.540***
	(0.129)	(0.169)	(0.179)	(0.169)	(0.168)	(0.208)	(2.164)	(2.151)
Lower trade-to-GDP ratio	10.120***	9.654***	8.086*		7.955**	12.781***	50.172	18.559*
	(3.544)	(3.604)	(4.442)		(3.675)	(3.621)	(66.158)	(11.143)
Trade salience × symmetry				3.598***				
				(0.953)				
Lower openness score					0.582**			
					(0.237)			
Allies × ln(time)		0.096***	0.102***	0.105***	0.097***	–0.374***	0.919***	0.916***
		(0.031)	(0.033)	(0.031)	(0.031)	(0.041)	(0.286)	(0.286)
Joint democracy × ln(time)		–0.149**	–0.131**	–0.134**	–0.149**	–0.663***	1.667*	1.663*
		(0.064)	(0.066)	(0.064)	(0.064)	(0.085)	(0.868)	(0.867)
Frailty[e]			0.134***				2.738***	2.737***
			(0.035)				(0.103)	(0.104)
Cases	1,450	1,450	1,450	1,484	1,450	1,450	1,450	1,450
Failures	1,439	1,439	1,439	1,473	1,439	1,439	1,439	1,439
Times at risk	241,561	241,561	241,561	263,232	241,561	241,561	241,561	241,561
Log-likelihood	–9,019.02	–9,012.09	–8,992.88	–9,252.33	–9,009.29	–8,026.97	–7,278.23	–7,277.01

[a] The table reports coefficients and standard errors. The asterisks indicate statistical significance based on two-tailed tests: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
[b] Duration dependence in Models 6–8 is parameterized as a cubic polynomial of time. The time coefficients and the constant are omitted to save space.
[c] The model uses the square root transformation of *lower trade-to-GDP ratio*.
[d] Replacing the alliance variable with UN voting similarity (Gartzke & Jo, 2002) does not change the substantive results, and the lower trade-to-GDP ratio remains significant at the 10% level.
[e] This is the logarithm of the estimated variance of the Gamma frailty.

Models 6–8, I complement the analysis through the use of the discrete-time PGM model (Jenkins, 2004–05).[31] The three reported discrete-time models include a cubic polynomial of time to account for duration dependence in a flexible manner. Model 6 replicates Model 2, and although the coefficients are not identical, no major differences can be found.[32] Model 7 includes an individual frailty parameter to account for unobserved heterogeneity; interdependence retains the correct sign, but its coefficient is smaller than the corresponding standard error. Even though Model 8, which replicates Model 7 with the square root transformation of the lower trade-to-GDP ratio, shows interdependence to be marginally significant, similarly to Model 3, Models 7–8 show that the effect of interdependence is not completely insensitive to the reported controls for unobserved heterogeneity. Unfortunately, improving the model specification to eliminate that heterogeneity might not be possible, as some of the omitted variables might be simply unobservable.

I finally estimate the aforementioned two-stage model of MID onset and duration to control for the potentially non-random selection of MID dyads. Even though the main interest is in the effect of interdependence on MID duration conditional on MID onset, a valid inference might still be precluded if an unobservable variable such as resolve is correlated with both interdependence and duration. If greater resolve is necessary to cross the MID threshold in more interdependent dyads, but greater resolve also prolongs MIDs, at the second stage of the process the negative relationship between interdependence and MID duration might appear weaker although not necessarily reversed, as Levy (2003) argues.

In the first stage, I practically replicate a model reported by Oneal & Russett (1999: 433), who use a logistic regression with peace year splines to estimate the effect of the lower dyadic trade-to-GDP ratio on MID onset probability in all dyads. The right-hand regressors in the first-stage specification complement the Model 2 covariates with minor-power status (1 = Yes, 0 = No), the higher trade-to-GDP ratio, which Oneal & Russett use to control for asymmetry, and four peace years splines with evenly spaced knots.[33] The second-stage specification is the same as Model 2, but adds the predicted probability of onset from the first stage together with its time interaction. The presence of the minor-power status and higher trade-to-GDP ratio variables in the first- but not in the second-stage model is to satisfy the exclusion restrictions. Although I do not have theoretical grounds to exclude those variables in particular, they seem good instruments because they are correlated with MID onset but not with duration.

I report bootstrapped confidence intervals at the second stage in order to account for the additional uncertainty induced by the use of a predicted variable from the first stage. Bootstrapping is a computationally intensive method, but it circumvents the necessity for complex analytical derivations. I perform 1,000 bootstrap replications, in each of which I first draw with replacement from the original sample a sample of equal size. I then estimate the first-stage logit, excluding dyadic years with ongoing MIDs but no new initiations. I then predict the

[31] The models were estimated with the *pgmhaz8* program for Stata, available from Jenkins (2004–05).

[32] The *alliance* and *joint democracy* coefficients differ in magnitude although their signs are the same. It seems that this difference is caused by the time interactions. When a complementary log-log model that produces almost identical estimates to the PGM (Jenkins, 2004–05) is estimated without the time interactions, the coefficients are almost identical to those in Model 2. Moreover, even with the time interactions, when MID day splines are substituted for the cubic polynomial of time, the *alliance* and *joint democracy* coefficients are again much closer to those in Model 2. These alternative specifications are reported in the accompanying log file.

[33] Generated using Tucker's *btscs* program for Stata (Beck, Katz & Tucker, 1998).

onset probability for each dyadic year, use it to estimate the second-stage model, and save the coefficients.[34] The bootstrapped second-stage coefficients are reported in Table II.

As Table II shows, the first-stage model is similar to the one estimated by Oneal & Russett (1999). The second-stage model does not indicate major changes in the covariates from Model 2, and, in fact, they appear with similar coefficients, which suggests that the non-random selection of MIDs does not distort them significantly.

That this selection is most likely non-random is evident from the negative coefficient of onset probability at the second stage. As MIDs become more probable, they become more prolonged, which actually fits Levy's (2003) argument. For example, higher resolve might make MIDs more likely, but also longer, which is why we tend to observe longer MIDs than if dyads crossed the MID threshold randomly. Moreover, the effect of onset probability declines with time, which seems plausible as the longer a MID lasts, the less relevant the factors that caused it should appear. Most importantly, however, controlling for the non-random selection into MIDs does not change the reported shortening effect of interdependence on MID spells, which lends support to the argument advanced in the theoretical section.[35]

Conclusion

The empirical findings presented here show an inverse relationship between interdependence and the duration of militarized conflict. Although novel, this result should not appear illogical, given the way many international relations scholars now think about the nature of such conflict. The expanding bargaining literature on war cited in the theoretical section implies that asymmetric information drives both the onset and termination of militarized disputes. Therefore, if dispute onset and termination are governed by a single variable, factors that influence that variable should have a discernible impact on both types of events.

However, the trade and conflict literature has focused predominantly on the pre-onset effects of interdependence while being less vocal about the nature of its post-onset role. In contrast, the argument developed here, which builds on the opportunity-cost explanation for the commercial peace, suggests that trade increases the price extracted from fighting states to eliminate the informational uncertainty that makes fighting rational both before and after conflict onset. If asymmetric information drives both the onset and termination of conflict, interdependence, which increases the price of that information, should lead to both fewer and shorter disputes.

The trade and conflict literature has already provided empirical support for the inverse relationship between interdependence and dispute onset. This study complements that literature by also providing evidence for the existence of a similar inverse relationship between interdependence and conflict duration, a relationship which is robust under a number of alternative model specifications and concept operationalizations. Moreover, the study offers further support for the opportunity-cost argument, as the latter empirical relationship follows logically from that argument. Indeed, the study's main finding does not necessarily undermine signaling theory, which is more ambiguous about the expected effect of interdependence on conflict duration. However, a novel result

[34] The data are in the non-directed dyad-year format with the only unusual thing being the additional dyad-years which record second or greater MID onsets in that year. The additional dyad-years are excluded from the first-stage logit, but onset probability is predicted for them as well. Thus, multiple MIDs spells within the same dyad-year share the same predicted onset probability at the second stage. I use Stata's *bsample* command to draw each bootstrap sample.

[35] One might wonder whether the unobserved heterogeneity indicated by Models 4, 7, and 8 can be eliminated by controlling for MID selection. While I find that likely, investigating the possibility seems too computationally intensive.

Table II. Two-Stage Model of MID Onset and Duration, 1950–92

Logistic Regression of MID Onset in Non-directed Dyads, 1950–92[a]

Independent variable	*Coefficient*	*2.5th percentile*	*97.5th percentile*
Ln(distance)	–0.507	–0.648	–0.366
Contiguity	2.604	2.263	2.946
Ln(power disparity)	–0.316	–0.407	–0.226
Minor powers	–2.572	–2.936	–2.208
Alliance	–0.351	–0.691	–0.012
Joint democracy	–1.118	–1.676	–0.561
Lower trade-to-GDP ratio	–56.960	–105.293	–8.629
Higher trade-to-GDP ratio	3.083	1.267	4.899
Peace years spline 1	0.011	0.009	0.013
Peace years spline 2	–0.012	–0.014	–0.009
Peace years spline 3	0.006	0.004	0.008
Peace years spline 4	–0.002	–0.003	–0.001
Constant	1.712	0.444	2.979
Observations	346,963		
Log-likelihood	–5,549.11		

Bootstrapped Cox Coefficients of Dyadic MID duration, 1950–92

Independent variable	*Coefficient*	*2.5th percentile*	*97.5th percentile*
Ln(distance)	0.072	0.017	0.126
Contiguity	0.274	0.106	0.454
Ln(power disparity)	0.036	–0.000	0.074
Alliance	–0.127	–0.355	0.084
Joint democracy	0.869	0.526	1.175
Lower trade-to-GDP ratio	9.456	4.330	29.571
Onset probability	–0.792	–1.589	–0.039
Alliance × ln(time)	0.108	0.051	0.169
Joint democracy × ln(time)	–0.136	–0.244	–0.014
Onset probability × ln(time)	0.116	–0.044	0.281
Bootstrap replications	1,000		

[a] The first-stage confidence intervals are based on robust standard errors clustered by dyad.

such as this should still stimulate further efforts to increase the precision of the competing theoretical explanations for the commercial peace until they can be differentiated empirically, even if that is not immediately possible.

In brief, previously it has been argued and empirically demonstrated that interdependence decreases the likelihood of conflict onset, but the literature has been less concerned with what happens during actual conflict. I provide evidence that the prospective opportunity costs of interrupted trade are also a powerful incentive for conflict termination as they increase the pressure on states to restore their normal relations, which would make commerce possible. Trade contributes to peace even after militarized conflict has already started by decreasing its duration, a finding implying that the pacifying effect of trade has so far been underestimated.

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