Online Appendix to “The Spread of Secrecy: Covert Military Alliances and Portfolio Consistency”

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This online appendix for “The Spread of Secrecy: Covert Military Alliances and Portfolio Consistency” contains:

1. An explanation for the paper’s choice of unit of analysis;
2. Evidence for the claim that partial and full secrecy should be treated synonymously in the statistical analysis;
3. A description of Klier and McMillen (2005)’s spatial logit model;
4. The calculations underlying Altonji et al’s robustness check for the strength of omitted variables;
5. Additional statistical test: A two-stage test for selection into alliances;
6. Additional statistical test: The more public alliances you have, the less you will create a secret one;
7. Additional statistical test: Joining a subsequent public alliance increases secret alliance failure; and
8. Additional statistical test: A hegemon’s secret pact induces secrecy in other states’ military alliances.

1 Justification for Unit of Analysis

This project uses dyads within an alliance instead of the alliance overall as the unit of analysis for several reasons. First, Fordham and Poast (2014) argue that alliances are the product of multilateral relationships and dynamics. Treating an alliance as a unified whole ignores how overlapping bilateral and multilateral dynamics converge in allowing states to create an alliance. For example, Poast (2010) points out how limited capabilities and geographic separation make Belgium and Turkey unlikely allies. However, their alliance "arguably has more to do with their relations with the United States than with each other." (Fordham and Poast, 2014) Moreover, certain members and relationships within a particular alliance have greater weight than others. The U.S. and the U.S.-Germany relationship are much more important to NATO's functioning than, say, the U.S. and Holland, particularly when it comes to outcomes like deterrence and burden sharing. Consequently, understanding alliance formation, design, and effects therefore requires that scholars choose a unit of analysis that reflects these dynamics as closely as possible, which in this case would be the security dyad.

Second, many alliances have asymmetric obligations. That is, members have different responsibilities, contribute varying types and amounts of support, and some may even have different foundational military commitments. As an example, the provisions of the U.S.-Japan security treaty concentrate exclusively on possible threats in East Asia. Japan is under no formal obligation to assist the U.S. in the event of an attack in the Western Hemisphere. Washington is granted use of military facilities in Japan, but Tokyo does not receive a reciprocal obligation in the U.S. Dyads or k-ads allow us to directly model these asymmetric obligations for each unit. If there were a hidden partner to an pact, for example, using the alliance as the unit of analysis would incorrectly code this relationship.

Third and extending from that, were alliances to serve as the unit of observation, there is a broader question of how to aggregate variables in a way that makes interpretive sense. Although democracies disproportionately ally with one another and are less likely to face military challenges, does the "average democracy" score of an alliance (or the aggregate, or the minimum, etc.) adequately capture those dynamics? (Siverson and King, 1980; Fearon, 1994; Lake, 1992; Reiter and Stam, 1998; Choi, 2001) Similarly, if a single alliance member is threatened with war, should its values for particular variables be used or that of the alliance overall? And what would it substantively mean to say that a high average democracy score makes an alliance (as opposed to a state) less likely to be challenged? Of course, dyads suffer from some of these considerations as well. But by using them as the unit of analysis, we can preserve greater amounts of information in available datasets, and we can leverage techniques like spatial econometrics to directly model how multilateral interdependence affects outcomes of interest.

Fourth and finally, when the international system possesses large multinational pacts, using alliances as the unit of analysis will significantly undercount the system's total number of security relationships, overinflating the importance of smaller, bilateral ties. To take simply one example, in 2003 (the last year for the ATOP3.0 dataset), 89.64 percent of alliances (173/193) were bilateral. But only 11.87 percent of *security dyads* (177/1499) were bilateral, because most countries now participate in large, multilateral partnerships like NATO. As important as they are, it would be odd if our choice of unit of analysis consistently gives the same weight to the U.S.-Japan and China-DPRK pacts as it does to the OAS, SCO, and NATO. Moreover, if we want a comprehensive model of the interstate security network rather than the inter-alliance network, then dyads and k-ads better reflect the former.

2 Partial and Complete Secrecy

In the empirical tests, this paper treats fully and partially secret security partnerships as analytically equivalent. Drawing from Ritter (2004), in both cases, the key issue is less whether a secret agreement or hidden provision exists, but what exactly it says. The net effect on an adversary's behavior is the same: caution regarding a target state and an improvement to deterrence. Yet, we may have strong reasons to think that even a partially public military pact grants members such benefits as a clear deterrent posture, defined policy boundaries, and delineated consequences if other states violate the agreement's terms. In that case, the fact that the statistical analysis finds no substantive difference between partial and completely secret alliances is evidence against the theory.

Fortunately, there are only 13 partially secret alliances in ATOP, and the text for some are in Gibler (2008). We can unpack this data to see which approach (i.e. partial and full secrecy are equivalent vs. they are not) better fits these cases. But first, it is critical to acknowledge discrepancies or missing information between ATOP's quantitative tables and its codesheets. The *former* include more detail on what provisions were secret versus public. Although Gibler (2008) fills in some of these details, not all partially secret ATOP alliances are contained in that volume. In addition, there may be some coding problems, where lack of data about an alliance was treated as secrecy, but it is difficult to make a clear judgment based on the codesheets.

Nevertheless, to distinguish between the two approaches to partial secrecy, we can examine two questions. The first is: What military obligations are being made public? Is the public portion akin to an "alliance," explicitly committing members to military action under conditions that may be kept secret? Or does the public document more closely resemble an entente or a friendship treaty, where members demonstrate common interest in particular policy areas? The former would support separating partially secret alliances into a separate category or even treating them as fully public, as allies gain deterrent benefits. The latter still leaves third-parties uncertain of what steps the putative partners might take, and indeed, may not even indicate that they are military partners.

As far as can be determined, the public portion of the partially secret alliances are always declarations of common interest, typically to cooperate to promote stability within a certain region, and commit signatories only to consult. The secret portions define "active" military obligations (like defensive/offensive commitments), as well as identifying the specific target, the conditions for invocation, each member's contribution, and the division of spoils.[1] For example, ATOP1080 was a consultative pact between Austria-Hungary and Russia from 1833-1859 declaring their interest in supporting the Ottoman Empire. The secret portion stated that their specific interest was curbing the Pasha of Egypt's autonomy and power. As conservative monarchies, these states had relatively friendly relations, as well as existing, overlapping defensive commitments (i.e. the Holy Alliance, the Quintuple Alliance). Similarly, Serbia and

Montenegro concluded ATOP1480 in advance of World War 1. The public portion included a declaration of friendship and consultation, while the covert part listed defensive and possibly offensive action against Turkey. Finally, ATOP2660 between France and Poland was also publicly consultative, but secretly defended against Germany during the Interwar period.

In all these alliances, members declare only anodyne statements of mutual interest and almost no policy boundaries. Granted, this does demonstrate friendship between members, so there is some signaling benefit. But such statements only establish the theory's prerequisite condition: positive relations. The more theoretically prominent and "direct" benefits of a public alliance – deterrence, open promises of military assistance, etc. – are likely not generated by these agreements, and indeed, Leeds, Long and Mitchell (2000) question whether pacts of this type ought to be classified as alliances at all. Instead, aligning with Ritter (2004), any deterrent benefits are "indirectly" generated, in that a simple statement of friendship and policy interest is unusual and it creates uncertainty about whether there are other, undisclosed military promises.

Indeed, compare these promises to public military obligations like NATO's Article 5 or the Quintuple Alliance's (1815) guarantees. The latter are categorical and cover a wide region. NATO commits members to defense against any threat to trans-Atlantic security. Quintuple Alliance members made similar promises for Continental Europe. While specifics on invocation, member contributions, and responsibilities are left out of the charter and adversaries certainly push the boundaries of what actions will invite an allied response, there is little question that that third parties can observe these military commitments, the "upper bound" of alliance action, and the geographic zone where it applies. By contrast, even the public portions of partially secret alliances are far more ambiguous about what members commit to, where, and with what limits. In that, they operate in a much more similar fashion to fully secret ones among friendly nations.

More briefly for the second question, do we see evidence of cumulativity? To some extent, yes. In six of the 13 partially secret pacts, members create additional covert alliances. Note that many of these states are peripheral to the international security system. These include the Kingdom of Two Sicilies in ATOP1025; Serbia and Montenegro in ATOP1480; the Malagasy Republic in ATOP3390; and four small Gulf Arab states in ATOP4965. As such, they have likely have fewer alliance options, public or secret, limiting the possible amount of cumulativity. As a final observation, none of the partially secret pacts fall in the 1873–1916 time period that defines the paper's central puzzle. This suggests that, while norms are an important consideration, portfolio consistency is a general mechanism explaining secrecy.

3 Klier and McMillen (2005) Spatial Logit Model

Klier et al's estimation equation for the spatial logit model is:

$$y^* = \rho W y^* + X\beta + \epsilon \tag{1}$$

$$\epsilon \sim \text{MVN}[0, \text{I}] \tag{2}$$

$$y_i = \begin{cases} 1 & \text{if } y_i^* > 0; \\ 0 & \text{if } y_i^* <= 0. \end{cases} \tag{3}$$

where X is a matrix of covariates, β is a vector of coefficients, and ϵ is an error term. The key component is the spatial lag term, $\rho W y^*$, where ρ is the coefficient for this term and y^* is the dependent variable. W is an NxN row-standardized square matrix, recording the relationship between the units of observation. The W matrix tracks the degree of connection across dyads, whether because they share state members or are part of the same security partnership. When incorporated into the model, W should control for these unit interdependencies. In addition, ρ, the coefficient on this matrix, measures the level of interdependence. Using spatial econometrics substitutes for checks using dyadic- or alliance-fixed effects, as this approach more directly accounts for unit interdependence. In that way, it also accounts in part for time, capturing time-invariant features of the dyadic relationship.

4 Altonji, Elder and Taber (2005) Robustness Check for Strength of Omitted Variables

To formalize their process, Altonji et al state the following condition:

$$\frac{\mathbb{E}(\epsilon|\text{Predictor}=1)-\mathbb{E}(\epsilon|\text{Predictor}=0)}{\text{Var}(\epsilon)}=\frac{\mathbb{E}(X'\gamma|\text{Predictor}=1)-\mathbb{E}(X'\gamma|\text{Predictor}=0)}{\text{Var}(X'\gamma)} \tag{4}$$

where X is the matrix of control variables for the outcome equation, γ is a vector of their coefficients, and ϵ is a vector of the residuals from the unobservables. In essence, on the left hand side, we calculate the potential effect that unobserved covariates could have on alliance participation, normalizing that for variation in the error term. On the right hand side, we do the same thing, normalizing for variation in our observed covariates. When this equality holds, a normalized shift in the distribution of unobservables would be equally as powerful as a shift in observables. Altonji et al then transform Equation 4 to ask how large the left hand side must be to explain away our predictor's effects, producing the following ratio, where β is our predictor estimate (in this case, the coefficient of *Secret Count*):

$$\frac{\hat{\beta}}{[\text{Var}(\text{Alliance})/\text{Var}(\text{Residuals})][\mathbb{E}(\epsilon|\text{Alliance}=1)-\mathbb{E}(\epsilon|\text{Alliance}=0)]} \tag{5}$$

Running this calculation on Model 4 in Table 1 provides the 20.41 ratio described in the article text.

5 Two-Stage Test for Selection

As a variant of omitted variable bias, the Altonji, Elder and Taber (2005) process just described should account for selection bias. But we can make an additional check against selection effects by using a two-stage process first modeling the decision to create an alliance, then the decision to make it secret. Doing so should separate the incentives to join an alliance from influencing those to make it secret.

I use the k-adic process discussed in the main text to construct a dataset of "shadow" non-allied cases so we have variation on the first dependent variable. Poast (2010) discusses why this is an effective method for modeling selection. Using all controls in both equations, Table I below presents the second-stage's results. As before, *Secret Count* is positively associated with *Secret*, further supporting the cumulative nature of covert pacts.

6 Previous Public Alliances and Secrecy

The main article demonstrates that secrecy begets secrecy. But what effect do prior public alliances have on the propensity to create a subsequent secret pact? The theory would expect that the more public alliances a country participates in, the less it will adopt a secret one, as covert partners will find their hidden status a signal of lower rank. I define *Public Count* analogously to *Secret Count*: a count of the number

Table I: Two-Stage Regression Results on *Secret Count* on Subsequent Alliance Secrecy.

	Second-Stage Results
Intercept	-3.86^{*}
	(0.63)
Secret Count	0.83^{*}
	(0.13)
Ally	1.24^{*}
	(0.28)
Open	-0.83^{*}
	(0.32)
CINC	-12.26^{*}
	(5.14)
Major	1.46^{*}
	(0.60)
Energy	-2.46×10^{-6}
	(3.56×10^{-6})
Production	2.72×10^{-5}
	(4.89×10^{-5})
Defect	0.54^{*}
	(0.17)
MID	-0.11
	(0.06)
Rival	0.14
	(0.13)
IGO	-0.05^{*}
	(0.02)
N	1555
AIC	190.48
BIC	447.24
$\log L$	-47.24

* indicates significance at $p < 0.05$

of public alliances a dyad participates in prior to creating a new security pact. In unreported models, I include this variable as a control in the main paper's models. It does not affect *Secret Count's* significance nor sign. In addition, *Public Count* has the expected negative sign and significance in those models.

More importantly, we may want to obtain precise estimates of this variable's effects. Model 1 in Table II below provides a baseline model using observed data. In Model 2, I run the matching algorithm on that data for *Public Count*.[2] I then rerun the logit regression. Finally, Model 3 applies the spatial logit approach on the matched data to account for unit interdependence. In all three, *Public Count* is negatively and significantly associated with *Secret*. On average, it reduces participation in a subsequent secret pact by 25.4 percent.[3]

7 Public Alliances and Secret Alliance Failure

An additional theoretical implication is that when a state joins a public pact, it should disrupt any existing secret alliances. To test this implication, I fashion two variables from the ATOP data and applied them to an additional statistical test. The dependent variable is *Secret Fail*. It takes a value of 1 if TERMCAUS in the ATOP dataset equals 6 or 8, 0 otherwise. TERMCAUS is a member-level variable that "offers our judgment regarding why an alliance member terminated its affiliation with a given alliance."[4] A 6 means that the alliance collapsed due to members engaging in military conflict with one another. An 8 indicates that a member violated an alliance provision, resulting in the pact's termination. Compared to the other possible values, these two are the clearest indicators of an alliance's internal cohesion.[5]

Table II: Effect of Prior Public Alliances on Likelihood of Creating a Secret Alliance.

	Model 1 Observed Data	Model 2 Matched Data	Model 3 Spatial Logit
Intercept	-4.60^*	7.10	5.78
	(0.74)	(4.02)	(5.95)
Public Count	-0.31^*	-0.74^*	-0.67^*
	(0.07)	(0.17)	(0.32)
Secret Count	0.56^*	0.16^*	0.13
	(0.05)	(0.08)	(0.11)
CINC	−5.57	14.93	11.96
	(3.11)	(7.87)	(11.65)
Major	1.75^*	0.70	0.76
	(0.39)	(0.70)	(0.82)
Open	−0.48	-1.82^*	−1.63
	(0.25)	(0.72)	(1.26)
MID	-0.24^*	-0.29^*	−0.25
	(0.07)	(0.11)	(0.18)
Rival	0.27	0.01	0.07
	(0.33)	(0.48)	(0.63)
IGO	0.00	0.15^*	0.13
	(0.01)	(0.07)	(0.11)
Defect	0.36	-4.19^*	−4.11
	(0.89)	(2.06)	(3.17)
Energy	0.10	−0.47	−0.38
	(0.11)	(0.25)	(0.44)
Production	0.01	−0.11	−0.13
	(0.09)	(0.20)	(0.25)
ρ			0.04
			(0.36)
N	3963	156	
AIC	380.93	124.26	
BIC	682.60	282.85	
$\log L$	−142.47	−10.13	

* indicates significance at $p < 0.05$

The independent variable is *Subsequent Public*, a dummy variable indicating whether a state joins a public alliance while participating in a previously-concluded covert pact. Although we can use the number of subsequent public agreements as our operationalization, the theory expects that *any* open pact will generate questions of rank for secret partners. Overall, it anticipates a positive and significant relationship between these variables.

That is what Table III displays. Model 1 uses observed data, while Model 2 uses matched data, which improves every covariate's balance. In both, subsequent public pacts increase the chance of secret alliance failure. Substantively, in Model 2, *Subsequent Public* leads to a 21.4 increase in secret alliance failure. In addition, subsequent public pacts *with major powers* should have an even larger substantive effect, as they often signal realignment and are frequently a secondary state's principal security partnership. This should more strongly affect the rank of covert pacts. Models 3 and 4 swap out *Subsequent Public* for *Subsequent Major Public*, and the estimated effects on *Secret Fail* are even larger. Substantively, it leads to a 31.2 percent increase in secret alliance failure. Overall, in line with the portfolio consistency mechanism, subsequent publicity undermines the rank of prior secrecy, leading to greater risk of alliance failure.

8 Hegemonic Secrecy

To buttress the case study, I ran additional statistical tests to determine whether the hegemon's participation in a covert alliance influences the secrecy/publicity of other military pacts. To do so, I include *Lead*, which indicates whether the leading

Table III: Effect of Subsequent Public Alliances on the Failure Rate of Prior Secret Alliances.

	Model 1	Model 2	Model 3	Model 4
Intercept	-6.29^{*}	-8.11^{*}	-5.99^{*}	-5.40^{*}
	(1.47)	(2.71)	(2.46)	(2.70)
Subsequent Public	2.52^{*}	2.80^{*}	3.13^{*}	2.74^{*}
	(0.79)	(1.01)	(0.1.02)	(1.01)
CINC	-18.84^{*}	-20.48	-18.78^{*}	-14.59
	(8.41)	(8.23)	(9.65)	(8.27)
Major	1.64^{*}	1.99^{*}	1.52^{*}	1.75^{*}
	(0.68)	(0.79)	(0.78)	(0.79)
Open	1.64^{*}	1.56^{*}	1.80^{*}	1.83^{*}
	(0.55)	(0.69)	(0.59)	(0.69)
MID	0.16^{*}	0.19^{*}	0.16^{*}	0.22^{*}
	(0.08)	(0.09)	(0.08)	(0.09)
Rival	1.06^{*}	1.06^{*}	1.26^{*}	0.71
	(0.50)	(0.57)	(0.54)	(0.57)
IGO	-0.03^{*}	-0.04^{*}	-0.03^{*}	-0.04^{*}
	(0.01)	(0.02)	(0.02)	(0.02)
Defect	-15.92	-16.92	-16.13	-15.96
	(1490.11)	(1298.01)	(2009.38)	(1306.67)
Energy	-0.67^{*}	-0.31	-0.77^{*}	-0.69
	(0.28)	(0.48)	(0.33)	(0.48)
Production	0.71^{*}	0.45	0.74^{*}	0.54
	(0.35)	(0.38)	(0.34)	(0.39)
N	4024	2967	4024	2196
AIC	126.34	105.66	122.66	104.15
BIC	403.54	354.54	399.87	354.13
$\log L$	-19.17	-8.01	-17.33	-7.80

* indicates significance at $p < 0.05$

state/hegemon is participating in a secret alliance that year, in the dataset and rerun the major models. If the theory is correct, the variable should be positively and significantly associated with *Secret*. Table IV below presents the results.

Lead has a positive coefficient and is statistically significant, supporting the claim that a central, hidden hegemonic alliance prompts secrecy among other alliances. These effects hold despite the corrections for imbalances in the data, unit interdependence, and k-adic adjustment. While the substantive effect attenuates across these robustness checks, *Lead* possesses the strongest substantive effect of all the significant and positive variables. (Although note that, because it is a dichotomous variable, *Lead's* effects may not be as strong as higher values of *Secret Count*.) *Secret Count* remains significant and with the expected sign. Interestingly, when significant, *Open* continues to be negatively associated with secrecy. In total, these results support the contention that the hegemon's secrecy induces secrecy in other pacts, buttressing the case study's findings.

Table IV: Statistical Results with *Secret* as Dependent Variable, Adding *Lead* as Explanatory Variable.

	Model 1 Raw Data	Model 2 Matched Data	Model 3 Spatial	Model 4 K-Adic
(Intercept)	-5.48^{*}	-4.30^{*}	-3.30^{*}	-3.77^{*}
	(0.70)	(0.97)	(0.83)	(0.57)
Secret Count	0.62^{*}	0.44^{*}	0.48^{*}	0.90^{*}
	(0.05)	(0.05)	(0.05)	(0.14)
Lead	3.00^{*}	2.70^{*}	2.44^{*}	2.03^{*}
	(0.53)	(0.54)	(0.52)	(0.58)
Open	-0.50^{*}	0.19	-0.27	-1.16^{*}
	(0.25)	(0.33)	(0.26)	(0.34)
MID	-0.14^{*}	-0.12	-0.14	-0.08
	(0.07)	(0.09)	(0.09)	(0.06)
Major	1.11^{*}	-1.82	-1.90	0.98
	(0.39)	(1.33)	(1.26)	(0.59)
CINC	-4.32	3.98	3.82	-10.78^{*}
	(3.51)	(5.18)	(4.91)	(5.26)
Rival	-0.10	-0.14	-0.18	0.07
	(0.35)	(0.32)	(0.32)	(0.14)
IGO	-0.00	0.02^{*}	0.02	-0.03
	(0.01)	(0.01)	(0.01)	(0.02)
Defect	-0.38	-1.44	-1.33	0.58^{*}
	(0.91)	(0.91)	(0.86)	(0.17)
Energy	0.06	-0.13	-0.12	-2.91×10^{-6}
	(0.11)	(0.12)	(0.12)	(3.72×10^{-6})
Production	-0.01	0.11	-0.08	3.65×10^{-5}
	(0.09)	(0.11)	(0.11)	(5.00×10^{-5})
N	3963	616	616	1559
AIC	384.31	289.52		199.86
BIC	685.98	537.22		456.75
$\log L$	-144.15	-88.76		-51.93

* indicates significance at $p < 0.05$

References

Altonji, Joseph, Todd Elder and Christopher Taber. 2005. “Selection on Observed and Unobserved Variables: Assessing the Effectiveness of Catholic Schools.” *Journal of Political Economy* 113(1):151–184.

Choi, Ajin. 2001. Democracy, Alliances, and War Performance in Militarized International Conflicts, 1816-1992 PhD thesis Duke University Durham, NC: .

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

Fordham, Benjamin and Paul Poast. 2014. “All Alliances are Multilateral: Rethinking Alliance Formation in an International System.” *Journal of Conflict Resolution* 58.

Gibler, D. 2008. *International Military Alliances, 1648-2008*. CQ Press.

Klier, Thomas H. and Dan McMillen. 2005. “Clustering of auto supplier plants in the U.S.: GMM spatial logit for large samples.” *Working Paper Series* WP-05-18. Federal Reserve Bank of Chicago.
URL: *”http://ideas.repec.org/p/fip/fedhwp/wp-05-18.html”*

Lake, David. 1992. “Powerful Pacifists: Democratic States and War.” *American Political Science Review* 86:24–37.

Leeds, Brett Ashley, Andrew Long and Sara Mitchell. 2000. “Reevaluating Alliance Reliability: Specific Threats, Specific Promises.” *The Journal of Conflict Resolution* 44(5):686–699.

Poast, Paul. 2010. “(Mis)Using Dyadic Data to Analyze Multilateral Events.” *Political Analysis* pp. 403–425.

Reiter, Dan and A. Stam. 1998. “Democracy and Battlefield Military Effectiveness.” *Journal of Conflict Resolution* 42:259–78.

Ritter, Jeffrey Munro. 2004. ”Silent Partners” and Other Essays on Alliance Politics PhD thesis Harvard University.

Siverson, Randolph M. and Joel King. 1980. “Attributes of national alliance membership and war participation, 1815-1965.” *American Journal of Political Science* 24:1–15.