Does Voting Order Matter? Sequential influence in United Nations roll-call voting *

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Abstract

I study how sequential voting shape voters' behavior in collective choice. Using United Nations General Assembly roll-call voting data, I exploit random variation in voting order to examine how countries follow after observing others' choices. The reduced-form evidence suggests countries voting later respond to earlier votes. The responses are driven not only by aggregate information but also bilateral relations when economically or politically close states vote before a given country. I further construct a sequential model and perform counterfactual to illustrate the impact of underlying mechanisms.

Introduction

Voting in large committees is often modeled as a simultaneous choice: each voter decides based on their private preferences and information, and the final vote count aggregates these preferences and information. In many settings, however, voting proceed sequentially and the decisions are observed by the later voters. Earlier voting decisions affect later voters, by revealing information about the underlying merits of a proposal, by exerting pressure through political alliances and bilateral relationships, or other potential channels. Distinguishing sequential influence from preference of the voters and studying the mechanisms through which sequential influence operates are difficult because voting order is often endogenous to institutional rules or strategic behaviors.

This paper studies sequential influence using roll-call votes in the United Nations General Assembly during the early Cold War. A distinctive feature of UN roll calls in this period is that the country voting first was selected by lot, then the countries proceeded alphabetically, generating quasi-random variation in the ordering. I combine historical roll-call records with country-level bilateral relationships to examine whether, and through which channels, earlier votes affect decisions of later voters. The reduced-form evidence suggests that countries' voting choices respond to earlier votes, with the direction and magnitude of these responses varying across earlier voters. In particular, greater alignment is observed when certain politically relevant or closely allied countries vote earlier. The results suggest that sequential voting can operate through political influence beyond possible learning.

To better illustrate the patterns, I develop a sequential voting model. Countries update their beliefs about merits of the proposal after observing earlier votes and also face relationship/alliance dependent incentives to agree or disagree with particular countries. The model

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separates informational learning from political pressure. I also perform a set of counterfactual analyses: compare observed roll-call voting with simultaneous voting that removes sequential influence, evaluate alternative ordering rules that minimize alliance-driven distortion relative to a benchmark preference, and assess the importance of key players by removing a powerful country's alliance ties and measuring the resulting change in vote outcomes.

Literature

This paper relates first to the literature on social learning and sequential decision-making. When previous actions are observed before an agent makes decision, earlier choices can reveal private information and affect the decisions of later agents. Banerjee (1992), for example, shows how rational inference from the predecessor behavior generates herding, such that individuals follow earlier actions even when their own private information disagrees. Hermalin (1998) studies leadership as an informational mechanism: an informed leader can convey information through an observable action, inducing others to follow. In my setting, earlier votes in a roll call can similarly reveal information to countries voting later in the sequence.

Several empirical papers bring these ideas to sequential committee decision-making. Knight and Schiff (2010) estimate a model of social learning in U.S. presidential primaries, where vote outcomes in early states reveal information to voters from states that vote later, and use the estimated model to study alternative primary arrangements. Spenkuch, Montagnes, and Magleby (2018) study sequential roll-call voting in the U.S. Senate and exploit alphabetical voting order to identify position-dependent behavior, distinguishing strategic backward induction from alternative channels such as common-value learning. López-Moctezuma (2016) develops a model of sequential deliberation in the Federal Open Market Committee in which members with heterogeneous preferences and information update their beliefs after observing their peers' recommendations; structural model is then used to quantify the informational value of deliberation. Related work by Newham and Midjord (2019) uses the transition from sequential to simultaneous voting in FDA advisory committees to study informational herding and compare information aggregation across voting procedures. More broadly, Iaryczower and Shum (2012) structurally separate ideological preferences from private information in judicial decision-making, while Iaryczower, Shi, and Shum (2018) study the role of pre-vote deliberation in U.S. appellate courts.

The paper also contributes to the extensive literature using United Nations voting to study international political alignment. Early work emphasized the importance of voting blocs and coordinated behavior within the General Assembly (Hovet, 1960), while subsequent research examined the broader political divisions from coalitions of states. Krasner (1985), in particular, emphasizes conflicts between industrialized and developing countries (the developed north vs the developing south), while Voeten (2000) uses UN General Assembly roll calls to recover the principal dimensions of political conflict and voting alignment. This literature establishes that UN votes contain substantial information about alliances, geopolitical blocs, and preferences of the states. Existing analyses focuses on country preferences in explaining the votes, and do not examine whether the within-roll-call sequence itself affects those votes. This paper exploits the sequential nature of the voting to study how countries respond to votes that they observe before making their own decisions.

Methodologically, the model presented in this paper is closely related to López-Moctezuma and Johnson (2025). They model the sequential voting procedure of the U.S. Supreme Court to study how justices learn from senior colleagues voting before them. The present paper applies a similar model to international voting, but the UN setting introduces an additional mechanism: predecessors differ not only in the information their votes may reveal, but also in their political

or alliance relationship with the country voting. This makes it possible to compare responses to the aggregate information contained in previous votes with responses to the identities and political affiliations of the countries casting them.

The paper therefore connects these literature by treating UN roll calls as a setting in which social learning and political influence operate simultaneously. The lottery determining the first country that votes generates quasi-random variation in whether particular states have already voted when a country makes its decision. I use this variation to estimate order effects[1] and examine whether they depend on the identity and political proximity with earlier voters. I then develop a sequential voting model intended that separates belief updating from alliance-dependent incentives.

Data

I use historical roll-call voting records and text data from United Nations meeting and general debates. UN voting records contain country-level votes. Text data records the voting data and allows me to extract countries' political positions and contents of the resolutions under consideration.

Roll-call voting data

The roll-call voting data is *United Nations Roll Call Data, 1946–1985* compiled by the Inter-university Consortium for Political and Social Research (ICPSR Study 5512) (Inter-university Consortium for Political and Social Research, 1992). It records country votes in the United Nations General Assembly, including plenary meetings, special sessions, and subcommittees. Each observation is a roll-call vote on a specific proposal, such as a draft resolution, an amendment to a draft resolution, an individual provision, or a procedural motion. Multiple votes may relate to the same agenda item. In below, I use roll call and vote interchangeably. For each roll call, the data identify the voting choices of participating member states and the voting body it belongs (subcommittee, plenary or special sessions).

I transform the voting records into a country–vote panel. For each vote, country i's vote choice are coded as Yes, No, or Abstain and the country names are harmonized across historical changes in membership and naming conventions.

Official UN meeting records

For the UN meeting text records, I download the English-language meeting documents from the United Nations Official Document System (ODS) [2]. The records include plenary verbatim records (A/PV), records for subcommittees and special-session proceedings.

Many of the historical documents are available as scanned PDFs rather than machine-readable text. I process the original documents using Google Cloud Document AI. The text data are cleaned so that speakers and voting countries in the meeting records can be matched to the ICPSR roll-call data.

The text data contains the substantive statements made by country representatives during the discussions of agenda items that were voted on. They were used as a candidate for recovering the voting order and the latter to construct text-based measures of the political and substantive dimensions underlying each resolution. The data also contains information on voting order which will be detailed below.

[1] I sometimes coin this as leading effects of certain country in sequential voting context of this paper.
[2] Source: https://documents.un.org/.

Reconstructing voting order

The UN meeting texts identify the country called first in each roll call voting sequence. Within my study period (early cold war), for a normal roll call, the first country is selected by lot, after which voting proceeds alphabetically until all countries have voted. The identity of the first voter therefore determines the realized ordering of the entire roll call.

I use the first country recorded in the meeting text together with the alphabetical ordering of the participating member states to reconstruct the complete voting sequence for each roll call. Specifically, I order countries alphabetically, begin the sequence at the country identified as voting first in the official record, and continue cyclically through the alphabetical list. This procedure yields the position of every participating country in the realized roll call/vote.

The reconstructed sequence allows me to determine, for each country–vote observation, which countries had already cast votes. For a reference country c, I define

$$\text{Before}_{it}^{(c)} = \mathbb{1}\{c \text{ is called before } i \text{ in roll call } t\}.$$

The empirical sample is restricted to roll calls for which the ICPSR voting record can be matched reliably to the official meeting text record and for which the first voter can be identified. The analysis is further limited to the pre-1970 period, when roll-call voting remained common in the General Assembly.[3] The realized voting order was publicly available to countries at the time they voted.

The final voting sample contains 3,032 roll calls across plenary, subcommittees, and special-session proceedings, with 236,385 votes cast. Among votes cast, 57.1 percent are Yes, 22.7 percent are No, and 20.2 percent are Abstain. The average roll call contains 78.0 participating voters. Special sessions account for 48 roll calls and 4,152 votes cast, with an average of 86.5 participating voters per roll call; their vote distribution is similar to that of the full sample, with 58.6 percent Yes, 20.6 percent No, and 20.8 percent Abstain.

Matching speeches to roll calls

I divide each meeting text into agenda-item and match to corresponding vote records. I further identify individual countries from the speaker labels.

The basic textual observation is therefore a country–(roll-call)vote document: $d = (i, t)$, consisting of the speeches made by country i in the debate associated with roll call t.

Approximately 91 percent of roll calls can be matched to corresponding text in the meeting records.

Text-based measures of preferences and issue content

I use the matched speech text to construct measures of country ideal points along (potentially multi-dimensional) political spectrum (e.g., east-west during the Cold-War; South-North conflict between the developing and developed economies) . First, I estimate latent country positions from speech by national representatives. These country ideal points summarize systematic differences in political language used by different countries. I also back out the substantive content of each voted item as a topic intensity measure. The ideal points and topic intensity varies along the country dimension and roll call (vote) dimension, respectively.

We thus have three objects at the roll-call level: (i) countries' observed Yes, No, or Abstain decisions; (ii) the voting order; and (iii) text-derived measures of country preferences and vote substantive content measures. These measures are mainly used for structural estimation.

[3] With the introduction and increasing use of electronic voting equipment, simultaneous voting became more prevalent and sequential voting records less consistently available.

Given potential concerns around voters responding to the speeches made by other voting countries, I alternatively use General Debate speech that took place before vote begins to measure country preferences, where the vote characteristics are replaced by coarser roll call/ vote description variables.

Bilateral relationship measures

I supplement the voting and meeting-record data with proxies for pre-existing bilateral relationships.

I use the Correlates of War Diplomatic Exchange Data Set (Bayer, 2006). These data record the level of diplomatic representation between states. This data series is observed at five-year intervals. For the later part of the sample, I supplement these data with the Diplometrics Diplomatic Representation Database (Moyer, Turner, and Meisel, 2021) that provides annual measures. Together, these data provide proxies for the intensity of formal diplomatic relationships between states.

I also use annual bilateral trade data from Correlates of War Trade Data Set (Barbieri, Keshk, and Pollins, 2009) and the IMF Direction of Trade Statistics (DOTS)(International Monetary Fund, 2025). From these, I construct trade intensity between states.

Finally, I use the ICOW Colonial History Data Set (Hensel, 2018). The data record countries' principal colonial rulers and the states from which they obtained independence. I construct former colonial ties indicator.

Country identities are harmonized across data. The resulting bilateral relationship measures are matched to the country pairs and years corresponding to each UN roll call. In the reduced-form, I test whether responses to an earlier country's vote are stronger when the two countries have closer relationships. In the structural model, These bilateral relationship measures allow the effect of previously observed votes to vary systematically across country pairs.

Reduced-form Findings

Regression Specification

For a reference country c (e.g., USA), estimate at the country–roll call level (i, t):

$$Y_{it}^{(c)} = \alpha_i + \gamma_t + \beta_0 \,\text{Before}_{it}^{(c)} + \varepsilon_{it}^{(c)} \tag{1}$$

where the outcome $Y_{it}^{(c)}$ is measured in two ways relative to the reference country's vote v_{ct}:

$$Y_{it}^{(c)} = \begin{cases} \mathbb{1}\{v_{it} = v_{ct}\}, & \text{(discrete: alignment indicator)} \\ -|v_{it} - v_{ct}|, & \text{(continuous: negative vote distance).} \end{cases}$$

α_i: country FE, γ_t: roll call FE. $\text{Before}_{it}^{(c)} = 1$ if c is called before i in the roll-call t. $Y_{it}^{(c)} = 1$ if i vote the same vote as c; the continuous outcome takes the negative absolute distance between the two votes (on the ordinal Yes/Abstain/No scale), and higher values indicate more agreement in both specifications.

For expositional convenience, I refer to β_0 as the leading effect of country c: it captures how much more closely other countries vote with c when c votes first. A positive leading effect indicates the reference country's vote has greater subsequent alignment. I do not yet assign the effect to to a particular mechanism.

Special and Emergency Sessions

Table 1: Voting along c: order effects across reference countries, special and emergency sessions

	Subsample estimates				Full-sample interactions		
c	Full	1st W.	2nd W.	3rd W.	Before	× 1st W.	× 2nd W.
Panel A: Discrete outcome $Y_{it}^{(c)} = \mathbb{1}\{v_{it} = v_{ct}\}$							
USA	0.0407***	0.0328	0.0268*	0.0412***	0.0370*	0.0367	−0.0610
	(0.0133)	(0.0328)	(0.0157)	(0.0142)	(0.0197)	(0.0582)	(0.0833)
GBR	0.0380**	0.0232	0.0223	0.0316	0.0399*	0.0086	−0.0512
	(0.0176)	(0.0362)	(0.0366)	(0.0326)	(0.0231)	(0.0400)	(0.1126)
FRA	−0.0064	−0.0716**	0.0533*	0.0204	−0.0034	−0.0775	0.1334
	(0.0167)	(0.0309)	(0.0290)	(0.0178)	(0.0187)	(0.0540)	(0.1060)
USSR	−0.0154	0.0455***	−0.0050	−0.0309***	−0.0335*	0.0534	0.0883
	(0.0119)	(0.0166)	(0.0700)	(0.0114)	(0.0199)	(0.0512)	(0.0790)
CHN	0.0288**	0.0120	0.0410	0.0272	0.0483**	−0.0486	−0.0949
	(0.0137)	(0.0422)	(0.0294)	(0.0174)	(0.0219)	(0.0652)	(0.1115)
Panel B: Continuous outcome $Y_{ir}^{(c)} = -\lvert v_{ir} - v_{cr}\rvert$							
USA	0.0652***	0.0279	−0.0055	0.0693***	0.0647**	0.0550	−0.1620
	(0.0182)	(0.0360)	(0.0619)	(0.0186)	(0.0256)	(0.0861)	(0.1666)
GBR	0.0624***	0.0314	−0.0183	0.0568***	0.0742**	−0.0043	−0.1498
	(0.0230)	(0.0385)	(0.0728)	(0.0189)	(0.0334)	(0.0651)	(0.1915)
FRA	−0.0142	−0.0755***	0.0358**	0.0255	−0.0171	−0.0697	0.1739
	(0.0255)	(0.0259)	(0.0142)	(0.0222)	(0.0306)	(0.0625)	(0.1937)
USSR	−0.0525***	−0.0022	−0.0091	−0.0521***	−0.0628***	−0.0113	0.1763
	(0.0145)	(0.0326)	(0.0764)	(0.0110)	(0.0218)	(0.0828)	(0.1485)
CHN	0.0441***	0.0015	−0.0055	0.0430**	0.0832***	−0.0963	−0.1946
	(0.0157)	(0.0372)	(0.0619)	(0.0202)	(0.0320)	(0.0836)	(0.1957)
Country FE		Yes				Yes	
roll call FE		Yes				Yes	
Clustered SE				Yes (two-way)			

Notes: Each row reports the order effect for a different reference country c. The Full, 1st W., 2nd W., and 3rd W. columns report estimates of

$$Y_{it}^{(c)} = \alpha_i + \gamma_t + \beta_0 \text{Before}_{it}^{(c)} + \varepsilon_{it}$$

on the indicated sample. The final three columns report a full-sample specification interacting $\text{Before}_{it}^{(c)}$ with First- and Second-World indicators. The reported Before coefficient is therefore the effect for the omitted Third-World group, while the × 1st W. and × 2nd W. coefficients report the corresponding differential effects. All specifications include country and roll call fixed effects. Two-way clustered standard errors are reported in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Special and emergency special sessions are particularly useful for examining sequential influence of a given country's vote. These sessions are convened to address urgent issues, where

examples include the creation of the UN Emergency Force (UNEF) at the first emergency special session in 1956 (United Nations General Assembly, 1956) and the apportionment of peacekeeping costs at the fourth special session in 1963 (United Nations General Assembly, 1963a,b). Voting in these sessions is salient; even the United States "must husband its influence so that it can gain votes where it needs them most" (Keohane, 1967, p. 232). I also focus on the five permanent members of the Security Council and use them as reference countries given their central position in the conflicts raised in these sessions. For instance, the 1963 financing dispute had its roots in "disagreements between the powers," with the Soviet Union and France refusing to pay for UNEF and United States pressing for sanctions under Article 19 (Padelford, 1965, p. 444; Tandon, 1967). Literature also indicates that states of lesser power tracked the votes of great powers. The United States "could influence the voting of its allies in Western Europe and in Latin America" (Tandon, 1967, p. 258), other countries withheld payments for after Moscow and Paris did, which were "out of similar reasons or sheer opportunism" (Padelford, 1963, p. 550), and small states were seen as voting out of "fear of antagonizing the great power to which they are linked" (Tandon, 1967, p. 270).

For the subgroup analysis, I divide countries into First-, Second-, and Third-World groups.[4] Table 1 shows substantial heterogeneity in leading effects across the permanent five. The clearest positive effects appear for the United States, the United Kingdom, and China. When the United States votes earlier, the probability that a subsequent country votes the same increases by about 4.1 percentage points. There is also positive effect for the continuous vote proximity by 0.065. The corresponding estimates are similar for the United Kingdom, at 3.8 percentage points and 0.062. China also exhibits a positive leading effect in both specifications. These patterns are therefore not driven solely by a particular coding of voting agreement.

The effects are not uniform across major powers. France exhibits non significant leading effect, while the estimates for the USSR are negative, particularly for the continuous outcome. The interacted specifications broadly preserve these differences.

Leading Effects for Non-P5 Reference Countries

The leading effects documented above are not confined to the special sessions and permanent five. I run the baseline discrete-outcome regression in (1) using each non-P5 country as the reference country c. I also pool the available plenary, special-session, and subcommittee roll calls for this regression.

The countries are ordered by their raw coefficient. Many non-P5 estimates cluster around zero[5]. The positive estimates include non-aligned states and regional powers such as Indonesia (IDN) and Tunisia (TUN). From literature, non-aligned states acted as floor leaders in the Assembly (Keohane, 1967), and a substantial group of smaller powers stood between the great powers (Padelford, 1965). More broadly, regional and middle powers show leadership in multilateral settings (Nolte, 2010). The estimates are nonetheless heterogeneous: many estimates have confidence interval that includes zero and some are negative.

Additional mechanisms: predecessor votes and bilateral relationships

In this subsection, I examine possible mechanisms of the leading effect. As an earlier reference points out, countries follow votes of others "out of similar reasons or sheer opportunism"

[4]The Three Worlds classification groups Western-aligned developed countries as First World, the USSR and Soviet-aligned states as Second World, and the remaining developing and non-aligned countries as Third World.

[5]12 of the 121 estimates differ significantly from zero at the 5 percent level, compared with roughly 6 expected by chance if no country led.

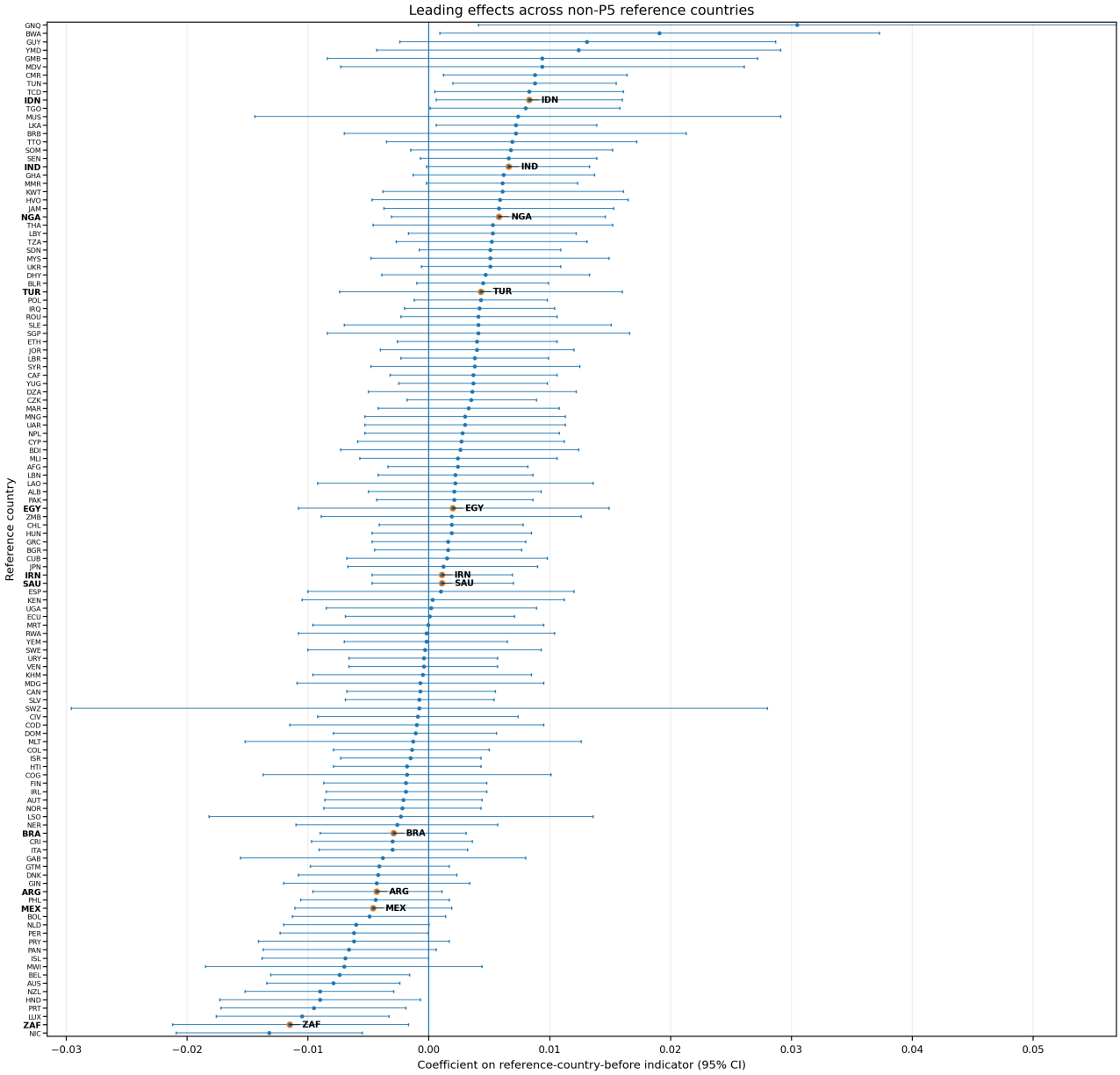


Figure 1: Leading effects across non-P5 reference countries

Notes: Each point reports the coefficient on $\text{Before}_{ir}^{(c)}$ from a separate country and resolution fixed-effects regression using country c as the reference country. The sample pools plenary, special-session, and subcommittee roll calls. Horizontal lines report 95% confidence intervals. Permanent Five countries are excluded.

(Padelford, 1963, p. 550). Earlier votes convey information on the substantive content and applicability of a proposal. The votes' political relevance may depend on who votes.

For each category $k \in \{\text{Yes}, \text{No}, \text{Abstain}\}$, I fit a linear probability regression with $Y_{it}^k = \mathbf{1}\{v_{it} = k\}$. Let H_{it} be the set of countries voting before i on roll call t, $n_{it} = |H_{it}|$, and N_t the total number of voters for each roll call. let $P_{ilt}^{(m)}$ denote the observed bilateral relationship measure of type m between focal country i and country l in the year of roll call t. For bilateral relationship measure m, define

$$a_{ilt}^{(m)} = P_{ilt}^{(m)} - \frac{1}{N_t - 1} \sum_{j \neq i} P_{ijt}^{(m)}. \tag{2}$$

The normalized exposures are

$$G_{it}^k = \frac{1}{n_{it}} \sum_{l \in H_{it}} \mathbf{1}\{v_{lt} = k\}, \tag{3}$$

$$\widetilde{B}_{it}^{k,(m)} = \frac{1}{n_{it}} \sum_{l \in H_{it}} a_{ilt}^{(m)} \mathbf{1}\{v_{lt} = k\}, \tag{4}$$

$$\widetilde{C}_{it}^{(m)} = \frac{1}{n_{it}} \sum_{l \in H_{it}} a_{ilt}^{(m)}. \tag{5}$$

All three denominators count every earlier actual voter. Thus G^k measures the predecessor share choosing k, $\widetilde{B}^{k,(m)}$ captures the additional exposure associated with relatively close predecessors (in terms of politcal or economic bilateral relation/ties) choosing k (normalized by the size of previous voters), and $\widetilde{C}^{(m)}$ controls for country i's relationship with the preceding voters (normalized by the size of previous voters). Countries that are first to vote are excluded from these share specifications. In general, B_{it} captures bilateral relation dependent incentives to follow, while G_{it} captures the remaining variation possibly attributable to learning.

The baseline equation is

$$Y_{it}^k = \alpha_i + \mathbf{z}_t'\gamma + \rho n_{it} + \beta_G G_{it}^k + \beta_B \widetilde{B}_{it}^{k,(m)} + \beta_C \widetilde{C}_{it}^{(m)} + u_{it}^k, \tag{6}$$

with coefficients estimated separately for each outcome and relationship measure. I include country fixed effects, voting venue (i.e., plenary meeting, special and emergency special sessions, subcommittees), and year indicators, and the number of earlier voters. These specifications do not include roll-call fixed effects as it kills much of the across resolution variation. Standard errors are two-way clustered by country and roll call.

Table 2 reports COW log trade and diplomatic presence, entered separately. Across all roll calls, both the general share (G_{it}) and relationship-weighted share (B_{it}) coefficients are positive for Yes, No and Abstain. For example, in the trade specifications the general share coefficients are 0.86, 0.81 and 0.63, respectively, while the relationship-weighted coefficients are 0.37, 0.33 and 0.26 (all $p < 0.001$). For a test of same regression on the subsample of special and emergency sessions[6], the general Yes and No associations remain positive, but the general abstention association is not statistically significant. The relationship-weighted associations remain positive for all three outcomes (all $p < 0.001$).

I next restrict the attention on the earlier voting countries to the permanent five. In the pooled P5 specification, the sums defining G^k, $\widetilde{B}^{k,(m)}$ and $\widetilde{C}^{(m)}$ run over $H_{it} \cap \mathcal{P}$, where $\mathcal{P}$ denotes the P5, but their denominators remain n_{it}, the number of all earlier voters. Centering also continues to use all other participants. Countries with predecessors but no earlier P5 voter

[6]the voting venue indicator are replaced by vote type indicator and year indicator replaced by session indicator

Table 2: Own-category predecessor shares and normalized relationship exposures

Sample / outcome / proxy	$\widehat{\beta}_G$	$\widehat{\beta}_B$	N
All / Yes / Trade	0.8643***	0.3654***	227,229
	(0.0322)	(0.0612)	
All / Yes / Dipl. presence	0.8567***	1.4102***	227,229
	(0.0312)	(0.1329)	
All / No / Trade	0.8146***	0.3307***	227,229
	(0.0289)	(0.0612)	
All / No / Dipl. presence	0.8101***	1.3912***	227,229
	(0.0284)	(0.1382)	
All / Abstain / Trade	0.6292***	0.2643***	227,229
	(0.0204)	(0.0462)	
All / Abstain / Dipl. presence	0.6349***	0.9379***	227,229
	(0.0202)	(0.0976)	
SS / Yes / Trade	0.5922***	0.5598***	4,008
	(0.1018)	(0.1027)	
SS / Yes / Dipl. presence	0.6475***	2.2065***	4,008
	(0.0996)	(0.2727)	
SS / No / Trade	0.5964***	0.8640***	4,008
	(0.0992)	(0.1461)	
SS / No / Dipl. presence	0.6525***	2.8143***	4,008
	(0.0968)	(0.3085)	
SS / Abstain / Trade	0.0352	0.3561***	4,008
	(0.1273)	(0.1003)	
SS / Abstain / Dipl. presence	0.0993	1.3219***	4,008
	(0.1231)	(0.3194)	

Notes: Two-way country/roll-call clustered standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$; unadjusted, two-sided tests. Country fixed effects are included; roll-call and position-bin fixed effects are not. Each row is a separate regression. All: source/year indicators; SS: session/vote-type indicators. Earlier-voter count and normalized composition $\widetilde{C}$ are included.

remain in the sample with zero P5 exposures, and P5 countries may themselves be focal voters. Relative vote position n_{it}/N_t are controlled for; for special sessions subsample, I control for the raw earlier-voter count.

Table 3 shows positive general P5 share associations for all three outcomes. Relationship-weighted Yes and No associations are positive for both proxies, while the abstention coefficients are not. The denominator makes these P5 exposures sensitive to the P5 share of the entire predecessor set, rather than averaging only among earlier P5 voters.

Table 3: Pooled P5 predecessor associations

Sample / outcome / proxy	$\widehat{\beta}_G$	$\widehat{\beta}_B$	N
All / Yes / Trade	1.0900***	0.3506***	227,229
	(0.3062)	(0.1295)	
All / Yes / Dipl. presence	1.2173***	1.1274**	227,229
	(0.2899)	(0.5520)	
All / No / Trade	1.2809***	0.3516**	227,229
	(0.3185)	(0.1371)	
All / No / Dipl. presence	1.4928***	1.0407*	227,229
	(0.3007)	(0.5411)	
All / Abstain / Trade	0.6305***	0.1529	227,229
	(0.1441)	(0.0948)	
All / Abstain / Dipl. presence	0.8520***	0.0318	227,229
	(0.1524)	(0.3352)	
SS / Yes / Trade	−0.3671	1.1640***	4,008
	(0.3790)	(0.2967)	
SS / Yes / Dipl. presence	0.1204	4.8360***	4,008
	(0.3645)	(1.2596)	
SS / No / Trade	−0.5405	1.4605***	4,008
	(0.8937)	(0.3854)	
SS / No / Dipl. presence	1.0881	3.5008**	4,008
	(0.9843)	(1.5388)	
SS / Abstain / Trade	−0.4872	0.3337	4,008
	(0.7723)	(0.2448)	
SS / Abstain / Dipl. presence	0.4439	−0.7311	4,008
	(0.5852)	(1.4216)	

Notes: Two-way country/roll-call clustered standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$; unadjusted, two-sided tests. Country fixed effects are included; roll-call fixed effects are not. Each row is a separate regression. All: source/year and relative-position controls. SS: session/vote-type and earlier-voter-count controls. Normalized P5 composition is included. All exposure denominators count all earlier actual voters.

To distinguish overlapping bilateral measures, I also include trade, average COW diplomatic representation level, and historical connections jointly, with a separate normalized composition control for each measure.

The historical indicator is the static union of recorded colonial-ruler and independence-from links; it is not a time-varying current-colony indicator. Table 4 shows positive conditional trade coefficients for all three outcomes in the all-source sample ($p = 0.014$, 0.037, and 0.032), while the diplomatic-level and historical coefficients are not statistically significant. In Special and emergency sessions subsample, trade remains positive for Yes and No and is less precisely estimated for Abstain ($p = 0.071$). The historical coefficients are negative for Yes and Abstain (both $p = 0.009$); the diplomatic-level coefficients are no statistically significant. By comparing the p5 only results to all precedessor country results, we can see an increase in the magnitude

of influence, consistent with the permanent five being more powerful in terms of their influence via bilateral political/economic ties, or being more knowledgeable about the vote.

Table 4: Pooled P5 associations with bilateral measures entered jointly

Sample / outcome	G	$\widetilde{B}$: trade	$\widetilde{B}$: dipl.	$\widetilde{B}$: history
All / Yes	1.1262^{***} (0.3158)	0.4212^{**} (0.1711)	-0.0637 (0.0897)	-0.4947 (1.0158)
All / No	1.2995^{***} (0.3274)	0.3474^{**} (0.1669)	0.0004 (0.0809)	0.0044 (0.8870)
All / Abstain	0.6915^{***} (0.1405)	0.2498^{**} (0.1167)	-0.0827 (0.0518)	-0.5016 (0.5096)
SS / Yes	-0.2899 (0.3764)	1.2758^{***} (0.3915)	0.2597 (0.1899)	-3.7791^{***} (1.4514)
SS / No	-0.6261 (1.0377)	1.4774^{***} (0.4453)	0.1396 (0.1866)	-2.2847 (2.5589)
SS / Abstain	-0.2454 (0.6553)	0.7246^{*} (0.4016)	-0.1021 (0.2089)	-2.6639^{***} (1.0133)

Notes: Two-way country/roll-call clustered standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$; unadjusted, two-sided tests. Country fixed effects are included; roll-call and position-bin fixed effects are not. All: N = 227,229, source/year and relative-position controls. SS: N = 4,008, session/vote-type and earlier-voter-count controls. Separate normalized composition terms for all three measures are included. Diplomacy is average representation level.

These results shows connections between earlier votes, bilateral relationships and subsequent choices. In particular, the positive coefficients on B indicates possible political influence and G captures the remaining variation attributable to learning.

Model

The reduced-form evidence above suggests that earlier votes may matter through more than one channel, but does not cleanly separate informational updating from country-pair-specific political influence. I develop a sequential voting model in which observed predecessor votes can both update a country's belief about the underlying state and affect its payoffs through directed political relationships. The model provides a framework for distinguishing these mechanisms and for studying how they interact with the realized voting order.

Primitives

A set of member states $i \in \{1, \ldots, N\}$ vote on a sequence of items indexed by t. On each item, states vote sequentially in a realized order, and $\mathcal{A}_{it}$ denotes the set of predecessors whose votes are publicly observed when it is country i's turn.

Each item has a latent binary state

$$\omega_t \in \{0, 1\},$$

where $\omega_t = 1$ indicates that supporting the proposal is substantively appropriate. Country i votes with posterior belief

$$q_{it} = \Pr(\omega_t = 1 \mid \Gamma_{it}), \tag{7}$$

where Γ_{it} contains its private information and the publicly observed voting history.

Country i has ideal point z_i. The substantive Yes and No positions on item t are denoted by x_t^Y and x_t^N, generating spatial losses

$$\ell_{it}^Y = \frac{1}{2}(z_i - x_t^Y)^2, \qquad \ell_{it}^N = \frac{1}{2}(z_i - x_t^N)^2. \tag{8}$$

A common stake $\varsigma > 0$ rewards taking the substantively appropriate position. In addition, earlier votes can affect country i through directed, alliance-sensitive political relationships. I distinguish two forms of such influence. Yes and No votes by predecessors can change the relative attractiveness of the two substantive positions, while predecessor abstentions can alter the relative attractiveness of abstaining.

The choice set is

$$v_{it} \in \{\text{No}, \text{Abstain}, \text{Yes}\}.$$

Payoffs

Yes and No combine spatial preferences, the appropriateness payoff, and pressure from predecessors who have taken a substantive position:

$$U_{it}(x_t^Y, \omega_t) = -\ell_{it}^Y + \frac{\varsigma}{2}\mathbf{1}\{\omega_t = 1\} + \frac{\varsigma}{2}S_{it}, \tag{9}$$

$$U_{it}(x_t^N, \omega_t) = -\ell_{it}^N + \frac{\varsigma}{2}\mathbf{1}\{\omega_t = 0\} - \frac{\varsigma}{2}S_{it}. \tag{10}$$

The signed alliance pressure term is

$$S_{it} = \sum_{l \in \mathcal{A}_{it}} \alpha_{il}^{\text{thr}} \, \text{dsign}(v_{lt}), \tag{11}$$

where

$$\text{dsign}(\text{Yes}, \text{No}, \text{Abstain}) = (+1, -1, 0).$$

Thus, for a positive link α_{il}^{thr}, an earlier Yes vote by country l raises the attractiveness of Yes relative to No, while an earlier No vote shifts incentives in the opposite direction. An abstention does not enter S_{it} because it does not take either substantive position.

Abstention has a separate, state-independent payoff. Let a_{it} denote its baseline net value, incorporating both the intrinsic value of avoiding a substantive position and any baseline cost associated with abstaining. Earlier abstentions may additionally affect this payoff through alliance-sensitive links:

$$U_{it}(A) = a_{it} + \frac{\varsigma}{2}R_{it}, \tag{12}$$

where

$$R_{it} = \sum_{l \in \mathcal{A}_{it}} \alpha_{il}^{d} \mathbf{1}\{v_{lt} = \text{Abstain}\}. \tag{13}$$

A positive α_{il}^d therefore makes abstention more attractive after country l has abstained. One interpretation is that an allied country's earlier abstention provides political cover for country i to avoid taking a decisive position. A negative coefficient instead makes abstention less attractive.

Importantly, abstention does not receive the appropriateness payoff, because it takes neither substantive side of the issue. Nor does it receive the signed Yes-versus-No pressure S_{it}. This does not imply that abstention is unaffected by influence from its allies: the separate term R_{it} allows predecessor abstentions to directly alter its payoff.

Expected utilities

Taking expectations over ω_t using posterior belief q_{it} gives

$$\mathbb{E}\left[U_{it}(x_t^Y) \mid \Gamma_{it}\right] = -\ell_{it}^Y + \frac{\varsigma}{2} q_{it} + \frac{\varsigma}{2} S_{it}, \tag{14}$$

$$\mathbb{E}\left[U_{it}(x_t^N) \mid \Gamma_{it}\right] = -\ell_{it}^N + \frac{\varsigma}{2}(1 - q_{it}) - \frac{\varsigma}{2} S_{it}, \tag{15}$$

$$u_{it}^A \equiv \mathbb{E}\left[U_{it}(A) \mid \Gamma_{it}\right] = a_{it} + \frac{\varsigma}{2} R_{it}. \tag{16}$$

The expected utilities from Yes and No vary with the country's posterior belief. The abstention payoff is independent of the latent state and therefore does not vary directly with q_{it}. It can, however, respond to the observed voting history through R_{it}.

The Yes–No comparison

The expected-utility difference between Yes and No is

$$\begin{aligned} \Delta_{it} &\equiv \mathbb{E}[U_{it}^Y] - \mathbb{E}[U_{it}^N] \\ &= \ell_{it}^N - \ell_{it}^Y + \varsigma \left(q_{it} - \frac{1}{2} \right) + \varsigma S_{it}. \end{aligned} \tag{17}$$

Using the spatial structure, this can be written as

$$\Delta_{it} = \varsigma \left[q_{it} - (1 - \widetilde{\pi}_{it}) \right], \tag{18}$$

where

$$1 - \widetilde{\pi}_{it} = 1 - \pi_{it} - S_{it}, \tag{19}$$

and

$$1 - \pi_{it} = \frac{1}{2} + \lambda_t(\kappa_t - z_i), \tag{20}$$

with

$$\kappa_t = \frac{x_t^N + x_t^Y}{2}, \qquad \lambda_t = \frac{x_t^Y - x_t^N}{\varsigma}. \tag{21}$$

Absent political pressure, $1 - \pi_{it}$ is therefore the posterior belief at which country i is indifferent between Yes and No. The observed substantive votes of predecessor countries shift this indifference threshold through S_{it}.

Abstention relative to a decisive choice

It is useful to define the average expected payoff from the two substantive positions:

$$\bar{U}_{it} \equiv \frac{1}{2} \left(\mathbb{E}[U_{it}^Y] + \mathbb{E}[U_{it}^N] \right). \tag{22}$$

Substituting (14) and (15),

$$\bar{U}_{it} = -\frac{1}{2} \left(\ell_{it}^Y + \ell_{it}^N \right) + \frac{\varsigma}{4}. \tag{23}$$

Both the posterior belief q_{it} and the signed political-pressure term S_{it} cancel from this expression.

Define the relative value of abstention as

$$b_{it} \equiv u_{it}^A - \bar{U}_{it}. \tag{24}$$

Using (16),

$$b_{it} = a_{it} - \bar{U}_{it} + \frac{\varsigma}{2} R_{it}. \tag{25}$$

Decision rule

Relative to $\bar{U}_{it}$,

$$\mathbb{E}[U_{it}^{Y}] - \bar{U}_{it} = \frac{1}{2}\Delta_{it}, \qquad \mathbb{E}[U_{it}^{N}] - \bar{U}_{it} = -\frac{1}{2}\Delta_{it}. \tag{26}$$

Therefore,

$$A \succeq Y \iff b_{it} \geq \frac{1}{2}\Delta_{it}, \tag{27}$$

$$A \succeq N \iff b_{it} \geq -\frac{1}{2}\Delta_{it}, \tag{28}$$

$$Y \succeq N \iff \Delta_{it} \geq 0. \tag{29}$$

When $b_{it} > 0$, abstention is optimal whenever

$$|\Delta_{it}| \leq 2b_{it}. \tag{30}$$

The resulting choice rule is

$$v_{it} = \begin{cases} \textsc{No}, & \Delta_{it} < -2b_{it}, \\ \textsc{Abstain}, & |\Delta_{it}| \leq 2b_{it}, \\ \textsc{Yes}, & \Delta_{it} > 2b_{it}. \end{cases} \tag{31}$$

If $b_{it} \leq 0$, abstention is dominated and the choice collapses to the binary Yes–No comparison

$$\Delta_{it} \gtrless 0.$$

Posterior decision thresholds

Substituting (18) into the decision rule gives

$$\textsc{Abstain} \iff |q_{it} - (1 - \widetilde{\pi}_{it})| \leq d_{it}, \tag{32}$$

where

$$d_{it} \equiv \frac{2b_{it}}{\varsigma}. \tag{33}$$

The corresponding posterior thresholds are

$$q_{it}^{N} = 1 - \widetilde{\pi}_{it} - d_{it}, \qquad q_{it}^{Y} = 1 - \widetilde{\pi}_{it} + d_{it}. \tag{34}$$

Hence,

$$v_{it} = \begin{cases} \textsc{No}, & q_{it} < q_{it}^{N}, \\ \textsc{Abstain}, & q_{it}^{N} \leq q_{it} \leq q_{it}^{Y}, \\ \textsc{Yes}, & q_{it} > q_{it}^{Y}. \end{cases} \tag{35}$$

The substantive votes of earlier countries affect $1 - \widetilde{\pi}_{it}$ through S_{it}, whereas earlier abstentions affect d_{it} through R_{it}.

Parameterizing the abstention payoff

Rather than separately parameterizing the gross benefit and baseline cost of abstaining, I parameterize their net contribution to the choice rule. Define the baseline abstention index

$$\theta_{d,it} \equiv \frac{2}{\varsigma}\left(a_{it} - \bar{U}_{it}\right). \tag{36}$$

Then

$$\frac{2b_{it}}{\varsigma} = \theta_{d,it} + R_{it}. \tag{37}$$

Because abstention should disappear when its relative value is non-positive, I use the smooth approximation

$$d_{it} = \frac{1}{\varrho}\log\left[1 + \exp\left\{\varrho\left(\theta_{d,it} + R_{it}\right)\right\}\right], \tag{38}$$

which satisfies

$$d_{it} \xrightarrow[\varrho\to\infty]{} \max\left\{\theta_{d,it} + R_{it}, 0\right\}. \tag{39}$$

In the empirical specification, the baseline abstention index is allowed to vary with resolution, country, and country–resolution covariates.

The parameter α_{il}^{d} in R_{it} measures how an earlier abstention by country l changes country i's willingness to abstain, in posterior-probability units.

Two channels of alliance-sensitive influence

The specification separates two ways in which bilateral country relationships can affect sequential voting.

First, substantive predecessor votes enter through

$$S_{it} = \sum_{l\in\mathcal{A}_{it}} \alpha_{il}^{\text{thr}}\text{dsign}(v_{lt}). \tag{40}$$

This channel affects whether country i favors Yes or No.

Second, predecessor abstentions enter through

$$R_{it} = \sum_{l\in\mathcal{A}_{it}} \alpha_{il}^{d}\mathbf{1}\{v_{lt} = \textsc{Abstain}\}. \tag{41}$$

This channel affects the attractiveness of abstaining relative to taking either substantive position.

The channels are separated by the observed action of the predecessor: a Yes or No vote enters S_{it}, while an abstention enters R_{it}. Thus, the model allows bilateral relationships to affect both the direction of a substantive vote and the willingness to avoid taking a substantive position, without forcing the two responses to operate through the same parameter.

Information and social learning

Before voting, country i observes a private signal about the latent state. Let

$$s_{it} \mid \omega_t \sim \mathcal{N}\left(\omega_t, \sigma_{it}^2\right), \tag{42}$$

and let

$$\rho_t = \Pr(\omega_t = 1)$$

denote the prior probability that supporting item t is appropriate.

Countries also observe the votes of predecessors. These actions are informative because they are generated by predecessors' private signals and preferences. Let L_{it}^{-} denote the public log-likelihood information accumulated from the observed voting history before country i votes, and let κ_L govern the weight placed on this social information.

Thus, the posterior q_{it} incorporates both the country's private signal and the information revealed by earlier public votes. This is the learning channel of the model. It is distinct from S_{it} and R_{it}: earlier votes affect beliefs through their informational content, while alliance-sensitive links allow the identity and position of the predecessor to directly affect utility.

From posterior thresholds to signal thresholds

For estimation, the posterior thresholds are mapped into the private-signal space in which the likelihood is evaluated. Inverting Bayes' rule gives

$$s_{it}^{v} = \frac{1}{2} + \sigma_{it}^{2}\left[\log\frac{q_{it}^{v}}{1-q_{it}^{v}} + \log\frac{1-\rho_t}{\rho_t} + \kappa_L L_{it}^{-}\right], \qquad v \in \{N, Y\}, \tag{43}$$

where q_{it}^{N} and q_{it}^{Y} are given by (34), and Appendix A provides the derivation.

Because the logit map is nonlinear, the abstention region that is symmetric around $1-\widetilde{\pi}_{it}$ in the posterior q_{it} becomes asymmetric in signal space. Its width there, $s_{it}^{Y} - s_{it}^{N}$, depends only on d_{it} and the signal scale σ_{it}^{2}: the prior ρ_t and the accumulated public information L_{it}^{-} enter both signal thresholds equally and therefore cancel from the width, shifting the region without resizing it.

The observed vote can therefore be written as a threshold rule in the private signal:

$$v_{it} = \begin{cases} \textsc{No}, & s_{it} < s_{it}^{N}, \\ \textsc{Abstain}, & s_{it}^{N} \le s_{it} \le s_{it}^{Y}, \\ \textsc{Yes}, & s_{it} > s_{it}^{Y}. \end{cases} \tag{44}$$

This representation yields the likelihood used in estimation. The realized history affects the voting rule through three conceptually distinct objects: public information from earlier votes, signed alliance-sensitive pressure from earlier Yes and No votes, and abstention-specific influence from earlier abstentions.

Graphical illustration

Figure 2 illustrates the three-choice decision problem.

The graphical language of a center and width is useful for visualizing the model: substantive predecessor votes move the location of the abstention region, whereas predecessor abstentions change how large that region is. In the formal model above, however, both effects arise directly from the underlying utilities and their corresponding posterior decision thresholds.

Identification

Random voting order, and learning versus alliance pressure

Our identification relies on the quasi-random order in which states vote. Order is assigned independently of a state's own preferences, so it changes the predecessor set $\mathcal{A}_{it}$— who has voted—

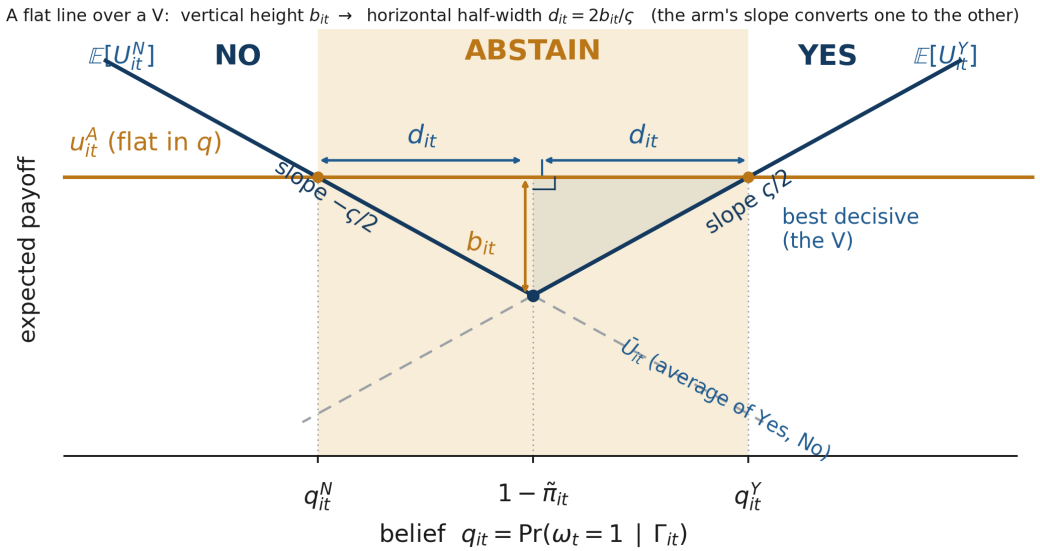


Figure 2: Expected payoffs from Yes, No, and Abstain as functions of the posterior belief $q_{it} = \Pr(\omega_t = 1 \mid \Gamma_{it})$. The expected payoffs from the two substantive positions are linear in q_{it}, while the abstention payoff is state-independent and therefore horizontal. Their intersections determine the posterior thresholds q_{it}^N and q_{it}^Y. Geometrically, $1 - \widetilde{\pi}_{it}$ is the center of the abstention region and $2d_{it}$ is its width. Earlier Yes and No votes shift the center through the signed political pressure S_{it}. Earlier abstentions change the width through R_{it} by altering the relative attractiveness of abstention. The figure therefore provides a graphical interpretation of the two distinct channels of alliance-sensitive influence.

for reasons unrelated to z_i or to i's private signal. A country that votes late systematically would otherwise face a particular set of predecessor countries; randomization severs that link.

Random order is also what separates learning from pressure stemming from political or other bilateral ties. When country i is preceded by a knowledgeable but unallied state—say a small, well-informed neighbor with no relationship to i—i's vote moves only through information; these observations isolate the learning weight. When i is preceded by a close ally whose information on the proposal that is already well known (i.e., similar signal precision with i), beliefs barely move but the pull from the preceding country's vote is strong which isolate the pressure. Because random order supplies both kinds of predecessor across roll calls, learning and pressure are separately identified.

If the pull depend only on the preceding country's own attributes, it would vary exactly as informativeness. In this way, the two channels could not be separated. Here, the pull depends also on the receiving country through the bilateral tie and the directed asymmetry between the two states. The same logic identifies the abstention-cover channel: an ally's abstention lifts i's willingness to abstain by an amount that depends on which ally abstains, so cover is separated from a state's own baseline tendency to abstain.

Separating preference from the value of abstaining

Kawai, Toyama, and Watanabe (2021) shows that if the same characteristics affects both how much a state cares turn out (choosing either yes or no) and how attractive it finds opting out (abstention), caring about the options is effectively equivalent to finding abstention more appealing. The margin in choice frequencies can not identify the turn out. The usual exclusion restriction assumes some covariates enters preferences but not the abstention margin. This work by ruling out the correlation between preference and participation which can be hard to

justify.

My model circumvents this via the structure of the payoffs (functional forms). For the two decisive outcomes, Yes rises with q_{it} exactly as No falls and meet at a benchmark $\bar{U}_{it}$. Averaging the frequencies of the two cancels the belief-dependent part and the antisymmetric alliance pressure. This leaves the benchmark quantity $\bar{U}_{it}$. Differencing yields the belief-dependent gap Δ_{it}. The Yes-vs-No margin locates the center $1 - \widetilde{\pi}_{it}$. How often a state abstains gives the height of the flat abstention payoff u^A_{it}, which is $b_{it} = u^A_{it} - \bar{U}_{it}$. This locates the width. As in Kawai, Toyama, and Watanabe (2021), the width is identified only as the net value of abstaining (i.e., net of cost), which is the $\theta_{d,it}$ here in my model.

Preferences and signal precision

A state's preference π_{it} and the roll call's prior ρ_t affects its leaning towards Yes and No; the signal precision σ_{it} scales both what it learns from earlier votes and whether it abstains. Following López-Moctezuma and Johnson, on each roll call the profile of votes is a mixture over the two possible states of the world ω_t; this pins the prior ρ_t and the probability of each vote given the state.

The preference bias π_{it} and the prior ρ_t are then separated by the pattern of variation each carries. The prior is common to every state on a given roll call which shifts the whole cross-section the same way. The bias varies with both the state and the roll cal. ρ_t changes moves every state in the same way for a given roll call; π_{it} changes show a a pattern that lines up with how those states vote elsewhere.

Signal precision σ_{it} governs how tightly its votes cluster around the choice it leans towards. A state with a precise signal tracks the merits closely and is little moved by earlier voters. A state with a imprecise signal follows the public record. We can leverage two sources of variation from the data. First source of variation is a given state's responsiveness to the accumulating (inferred) public information as countries vote before it. Second source of variation comes from the the first voters. Every state is chosen to vote first on some roll calls whose choice reflects only its own bias precision of its private signal, as well as the prior.

Estimation

I estimate the model in two stages. The first stage is the text model that estimates country ideal points and roll-call topic intensities that enter the structural covariates; the second stage is the full voting model estimation. The country-pair level covariates used to parameterize the bilateral realtion/alliance-pressure channels are the bilateral diplomatic, trade, and colonial-relationship measures described in the Data section.

Likelihood

Fix a roll call t with latent state $\omega_t \in \{0, 1\}$ and prior $\rho_t = \Pr(\omega_t = 1)$. Given $\omega_t = m$, country i's private signal is $s_{it} \mid \omega_t = m \sim \mathcal{N}(m, \sigma^2_{it})$, and the three-choice rule of the Model Section—No, Abstain, or Yes as s_{it} falls below s^N_{it}, between the cutoffs, or above s^Y_{it}—gives

state-contingent probabilities

$$P_{it}^{N}(m) = \Phi\left(\frac{s_{it}^{N} - m}{\sigma_{it}}\right), \tag{45}$$

$$P_{it}^{Y}(m) = 1 - \Phi\left(\frac{s_{it}^{Y} - m}{\sigma_{it}}\right), \tag{46}$$

$$P_{it}^{A}(m) = \Phi\left(\frac{s_{it}^{Y} - m}{\sigma_{it}}\right) - \Phi\left(\frac{s_{it}^{N} - m}{\sigma_{it}}\right), \tag{47}$$

where the signal cutoffs $s_{it}^{N} \leq s_{it}^{Y}$ are the Bayesian inversion of the posterior band, Eq. (43), and embed the pressure-adjusted centre $1 - \tilde{\pi}_{it}$, the wedge d_{it}, the prior ρ_t, and the learning statistic L_{it}^{-}.

The public statistic accumulates action-specific increments: each predecessor's realized vote is scored by its relative likelihood under the two states,

$$L_{it}^{-} = \sum_{l \in \mathcal{A}_{it}} \log \frac{P_{lt}^{v_{lt}}(0)}{P_{lt}^{v_{lt}}(1)}. \tag{48}$$

Because votes are public and cast in a known order, the joint probability of the observed votes on roll call t given $\omega_t = m$ factorizes along the order by the chain rule; marginalizing the common state yields the roll-call likelihood

$$p(v_{\cdot t} \mid \theta) = \sum_{m \in \{0,1\}} \rho_t^{m}(1 - \rho_t)^{1-m} \prod_{i} P_{it}^{v_{it}}(m), \tag{49}$$

with the product taken in voting order and each factor evaluated at the observed history. The sample log-likelihood is $\ell(\theta) = \sum_t \log p(v_{\cdot t} \mid \theta)$.

Learning conditional on alliance pressure

The increment (48) evaluates each predecessor's vote using the choice probabilities implied by that predecessor's pressure-adjusted cutpoints. Thus, the informational content of a vote is assessed conditional on the political incentives facing the predecessor, rather than interpreting those incentives themselves as information about ω_t. In particular, when political pressure makes a predecessor's action less sensitive to the underlying state, the resulting vote is less informative in the likelihood-ratio update. This allows the model to distinguish the informational content of earlier votes from their direct political effect on subsequent voters.

Maximum likelihood

I maximize $\ell(\theta)$ from multiple random starts and retain the best-converged optimum; standard errors are the square roots of the diagonal of the inverse observed information. The evidence in Appendix B confirms that this estimator recovers θ with negligible bias and near-nominal Wald coverage.

Alternative Bayesian estimation

As an alternative check, the structural model is also be estimated using Bayesian methods. I place weakly informative Gaussian priors on the coefficients entering the log- and logit-scale equations for σ_{it}, ρ_t, and π_{it}, and on the alliance and wedge coefficients. Posterior sampling uses Hamiltonian Monte Carlo with the no-U-turn sampler (Hoffman and Gelman, 2014).

Appendix B.3 assesses this implementation through simulation-based calibration. The rank statistics for all 31 parameters in the simulation specification are consistent with uniformity, providing evidence of computational calibration under that design. This exercise provides an additional check of the Bayesian implementation; the empirical results reported below use maximum likelihood with combined TBIP-posterior-draw and roll-call bootstrap inference.

Two-stage inference

The point estimates use a plug-in two-stage estimator. Country ideal points are the oriented variational posterior means from TBIP. Roll-call topic shares are constructed by normalizing posterior mean document intensities within each document and averaging these shares across documents associated with the roll call. The structural parameters maximize the likelihood evaluated at these text-derived covariates and the original voting sample.

To incorporate uncertainty in the text-derived covariates as well as roll-call sampling variation, I use a combined posterior-draw and roll-call bootstrap. In each replication, I draw country ideal points and document topic intensities from the fitted TBIP variational posterior, preserving the fitted coordinate orientation. I normalize the drawn document intensities, aggregate the resulting shares to roll calls, and reconstruct the associated structural covariates. I independently resample complete roll calls with replacement, retaining the ordered voting sequence within each selected roll call. The sampled votes remain matched to the corresponding drawn country positions, topic shares, and other covariates. A country's drawn position is shared across its appearances within a replication. I then re-estimate all structural parameters, using the original-sample fitted MLE as the initial value. The TBIP model itself is not re-estimated in each replication.

Let $\widehat{\boldsymbol{\psi}}$ denote the original-sample structural MLE and $\widehat{\boldsymbol{\psi}}^{(b)}$ the estimate from combined replication b. For parameter r, the bootstrap standard error is

$$\widehat{\mathrm{SE}}_{\mathrm{boot},r} = \left[\frac{1}{B-1} \sum_{b=1}^{B} \left(\widehat{\psi}_r^{(b)} - \overline{\psi}_r^{*} \right)^2 \right]^{1/2}, \qquad \overline{\psi}_r^{*} = \frac{1}{B} \sum_{b=1}^{B} \widehat{\psi}_r^{(b)}.$$

The reported 95% percentile intervals are the empirical 2.5th and 97.5th percentiles of the replicated estimates. I use $B = 500$ replications, all of which passed the numerical acceptance criteria. The bootstrap mean centers the standard-error calculation; the reported point estimate remains $\widehat{\boldsymbol{\psi}}$. Percentile intervals need not be symmetric around, or contain, that point estimate.

This hybrid procedure propagates uncertainty represented by the fitted TBIP variational posterior jointly with sampling variation across roll calls. Its interpretation relies on the adequacy of that posterior approximation and the roll-call resampling scheme; it does not account for error in the variational approximation itself.

Results

This section presents the results obtained so far.

Text-Based Model Estimates

To summarize countries' political positions and the substantive content of UN speeches and debates, I estimate a two-dimensional extension of the text-based ideal point (TBIP) model introduced by Vafa, Naidu, and Blei (2020) on the combined corpus. For document d and term

v,

$$y_{dv} \sim \text{Poisson}(\lambda_{dv}), \qquad \lambda_{dv} = \sum_{k=1}^{K} \theta_{dk}\beta_{kv} \exp\left(\mathbf{x}'_{i(d)}\boldsymbol{\eta}_{kv}\right), \tag{50}$$

where θ_{dk} is document d's intensity on topic k, β_{kv} is the baseline word weight for that topic, $\mathbf{x}_{i(d)} = (x_{1i(d)}, x_{2i(d)})$ is the two-dimensional ideal point of the country producing the document, and $\boldsymbol{\eta}_{kv}$ allows word usage within a topic to vary with political position. I estimate $K = 25$ topics jointly with the country ideal points using mean-field variational inference. This procedure approximates the posterior distribution with a tractable family of distributions and estimates their parameters by maximizing the evidence lower bound. For each country's two-dimensional ideal point, the variational distribution is Gaussian with an estimated location vector and dimension-specific scales.

The coordinates plotted in Figure 3 are the means of these approximate posterior distributions, after applying a common orthogonal transformation to orient the two-dimensional space. They therefore summarize estimated political positions rather than individual posterior draws. The orientation changes the coordinate representation while preserving distances between countries.

For the subsequent voting analysis, I aggregate the document-level topic intensities to the roll-call level. Let

$$\boldsymbol{\vartheta}_t = (\vartheta_{t1}, \ldots, \vartheta_{tK}) \tag{51}$$

denote the vector of normalized topic intensities for roll call t, obtained by averaging the corresponding document-level topic shares within the roll call. Thus the two text-based objects used below are the country position $\mathbf{x}_i = (x_{1i}, x_{2i})$ and the roll-call issue-content vector $\boldsymbol{\vartheta}_t$.

Figure 3 shows a clear separation between the Soviet bloc and much of the Western group along the first dimension, together with substantial additional cross-cutting variation along the second. I therefore use the two coordinates jointly as a flexible measure of political position rather than assigning a fixed substantive interpretation to either dimension.

The estimated topics also correspond closely to recognizable UN issues. Prominent topics include the Congo crisis, membership and Security Council questions, Chinese representation, Israel–Palestine, nuclear disarmament alongside colonial, human-rightsand economic-development topics.

These estimates provide the covariates used in the mechanism analysis. In particular, the benchmark vote prediction for country i on roll call t uses its two-dimensional position $\mathbf{x}_i$, the roll-call topic vector $\boldsymbol{\vartheta}_t$, and their interactions:

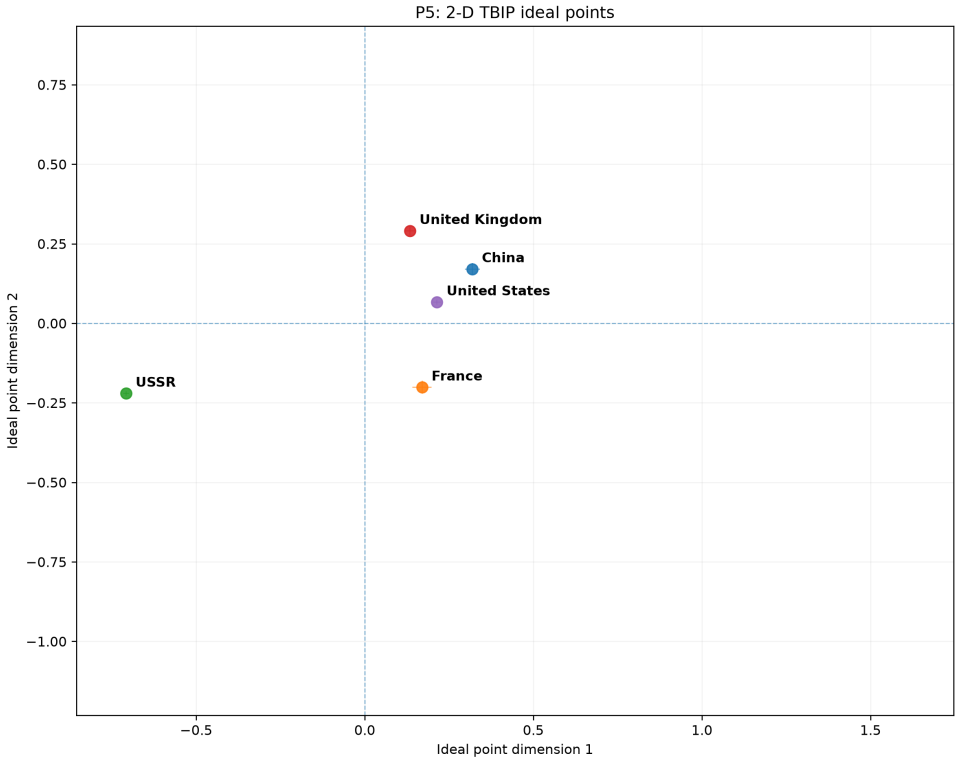
$$\mathbf{X}_{it} = [x_{1i},\, x_{2i},\, \boldsymbol{\vartheta}_t,\, x_{1i}\boldsymbol{\vartheta}_t,\, x_{2i}\boldsymbol{\vartheta}_t]\,. \tag{52}$$

This allows the predictable component of voting to vary jointly with persistent political position and the substantive content of the roll call.

Full Model Estimates

I next estimate the sequential model using the text-based measures of political position and topic intensities of the votes. The estimation sample contains 206,625 Yes, No, and Abstain decisions in 2,679 roll calls over 1948–1971.

I report structural maximum-likelihood point estimates together with standard errors and percentile intervals from a combined TBIP-posterior-draw and roll-call bootstrap. The point estimates use the original voting sample and mean-based text covariates. Country positions are the oriented variational posterior means from TBIP. For topic content, I normalize the



(a) Permanent members



(b) Countries by Three Worlds classification

Figure 3: Estimated two-dimensional text-based ideal points.

Notes: The Three Worlds classification groups Western-aligned developed countries as First World, the USSR and Soviet-aligned states as Second World, and the remaining developing and non-aligned countries as Third World.

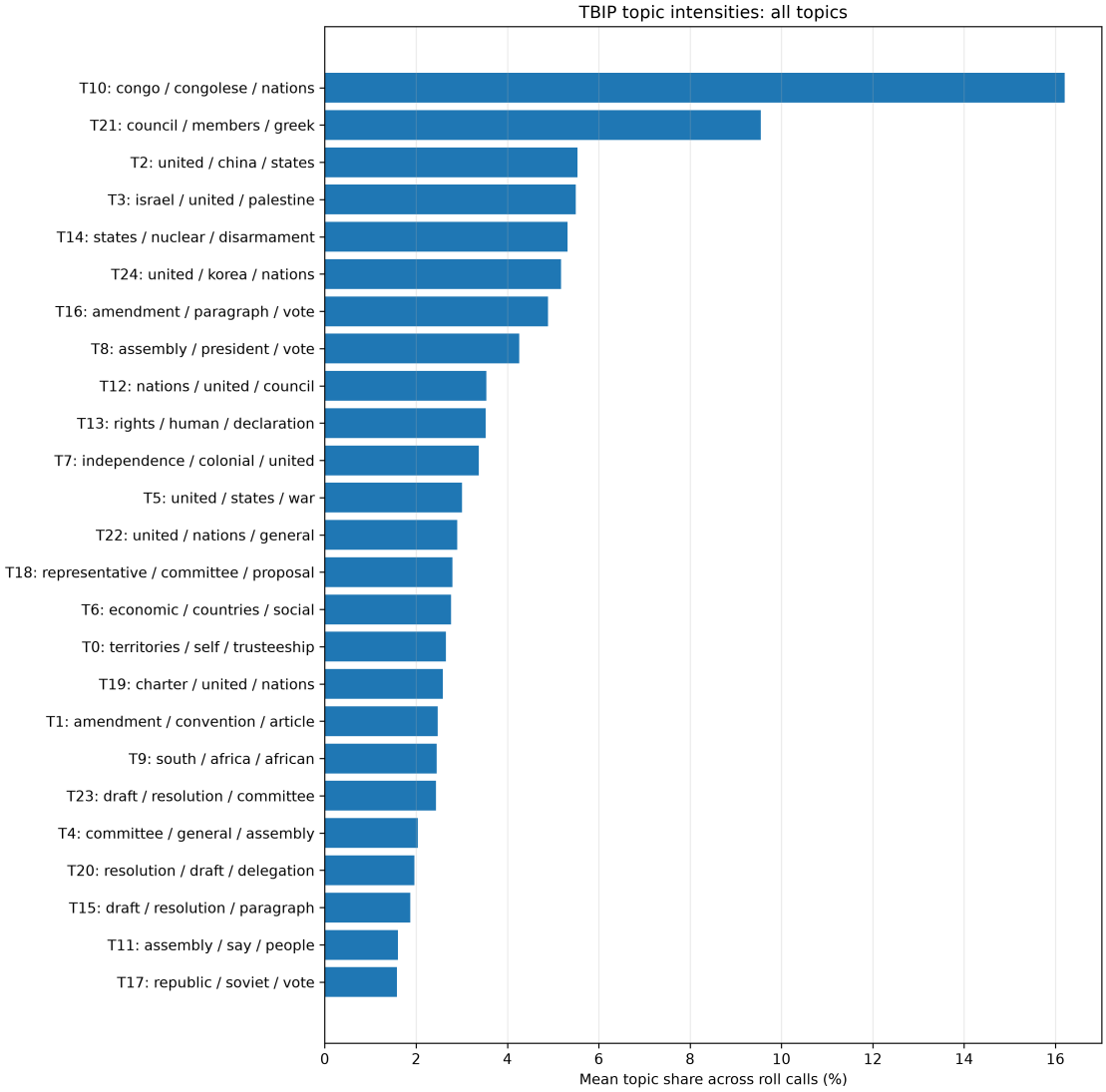


Figure 4: Estimated TBIP topic intensities. Bars report the mean normalized topic share across roll calls, after averaging document-level topic shares within each roll call.

variational posterior mean document intensities within each document and average the resulting shares within each roll call. Thus, for the documents $\mathcal{D}_t$ associated with roll call t,

$$\widehat{\vartheta}_{tk} = \frac{1}{|\mathcal{D}_t|} \sum_{d \in \mathcal{D}_t} \frac{E_q[\theta_{dk}]}{\sum_{h=1}^{25} E_q[\theta_{dh}]}.$$

This is a normalization of mean intensities, rather than the posterior mean of normalized shares. The structural parameters are estimated jointly conditional on these text-derived inputs.

Empirical covariates and parameterization. The empirical specification follows the resolution, country, and country–resolution block structure of Appendix B.2, while allowing the dimensions and included covariates to reflect the UN data. To distinguish these blocks from the TBIP country coordinates, let $\mathbf{X}_t = (1, C1_t, C3_t, H_t^z)^\top$, where $C1_t$ and $C3_t$ identify First Committee and Third Committee roll calls, respectively, and H_t^z is the standardized largest topic share, $H_t = \max_k \widehat{\vartheta}_{tk}$. The committee indicators proxy for security and humanitarian issue content; they are not text-based classifications of every roll call. Let

$$\mathbf{Z}_{it} = \left(P5_i,\, \text{TradePower}_{it}^z,\, \text{DiplomaticCapacity}_{it}^z\right)^\top,$$

and let W_{it} indicate that country i is named in the roll call's codebook description. Trade power is the standardized log of one plus total recorded trade with other sample entities. Diplomatic capacity is the standardized mean bilateral diplomatic-representation score across other sample entities. The information and prior equations are

$$\log \sigma_{it} = \mathbf{X}_t^\top \boldsymbol{\alpha}_\sigma + \mathbf{Z}_{it}^\top \boldsymbol{\beta}_\sigma + W_{it}\chi_\sigma, \tag{53}$$

$$\log \frac{\rho_t}{1-\rho_t} = \mathbf{X}_t^\top \boldsymbol{\delta}_\rho. \tag{54}$$

Since σ_{it} is a signal standard deviation, a positive coefficient in its equation corresponds to less precise private information.

Preferences depend directly on both TBIP coordinates, the first 24 of the 25 topic shares, and their interactions:

$$\log \frac{\pi_{it}}{1-\pi_{it}} = \zeta_{\pi,0} + \sum_{k=1}^{24} \zeta_{\pi,k} \widehat{\vartheta}_{tk} + \sum_{m=1}^{2} \eta_{\pi,m} \widehat{x}_{mi} + \sum_{m=1}^{2}\sum_{k=1}^{24} \lambda_{\pi,mk} \widehat{x}_{mi} \widehat{\vartheta}_{tk}. \tag{55}$$

Topic 25 is the omitted category. This specification retains the preference blocks used in the Monte Carlo exercise, but uses a logistic link for π_{it} rather than the additive specification for $1 - \pi_{it}$ in that exercise. All 75 preference coefficients are estimated within the structural likelihood; no preliminary vote regression is used to construct the preference index.

For the baseline abstention index I use

$$\theta_{d,it} = \mathbf{D}_t^\top \boldsymbol{\theta}_X, \qquad \mathbf{D}_t = (1, \text{Procedural}_t, [\log(1 + \text{Sponsors}_t)]^z)^\top. \tag{56}$$

Procedural status and the sponsor-count proxy are extracted from codebook wording; unrecognized sponsor counts are coded zero. Unlike the more general Monte Carlo specification, this application includes only resolution-level covariates in the baseline abstention index. The wedge also responds to predecessor abstentions through R_{it}, using (38) with $\varrho = 50$.

Both political-influence channels include five bilateral features and four P5 sender deviations. Let $A_{il} = \mathbf{1}\{l \in \mathcal{P}_5, i \notin \mathcal{P}_5\}$ and let $\mathbf{h}_{il,t}$ contain trade-power asymmetry, bilateral trade, diplomatic representation, and colonial ties. For $c \in \{\text{thr}, d\}$, I specify

$$\alpha_{il,t}^c = \mathbf{h}_{il,t}^\top \boldsymbol{\gamma}_c + A_{il} \left[b_c + \sum_{s \in \mathcal{P}_5 \setminus \{\text{USA}\}} d_{c,s} \mathbf{1}\{l = s\} \right]. \tag{57}$$

The USA is the reference sender, with $d_{c,\text{USA}} = 0$. b_c is its P5-to-non-P5 component, and $b_c + d_{c,s}$ is the corresponding component for sender s. Both components are zero when the recipient is itself a P5 member; other bilateral features can still produce influence between P5 countries. The extension adds eight parameters, four in each channel. The private-information equation continues to use a single P5 membership indicator.

For current voter i and a predecessor country l, the bilateral relation proxies include P5 indicator, trade related variables, diplomacy measure. Power asymmetries are measured since we have directed bilateral influence. Continuous pair features are divided by their standard deviations without centering, preserving zero exposure.

The bilateral measures are matched using the corrected country-code crosswalk. In this specification, the China voting identifier is linked to the ROC/Taiwan entity throughout the sample. We have constant ROC bilateral mapping; it does not separately model influence from the PRC outside the observed voting sequence.

Estimates and two-stage uncertainty. The resulting specification has 109 parameters. Tables 5 and 6 report the learning, information, prior, abstention, and bilateral-influence coefficients; the 75 jointly estimated preference coefficients are not displayed.

To construct bootstrap standard errors and confidence intervals, I perform 500 replications. In each replication, I draw country ideal points and document topic intensities from the fitted TBIP variational posterior. I normalize the drawn intensities of the documents, aggregate topic shares to roll calls, and reconstruct the text-derived structural covariates. I then sample roll calls with replacement, retaining each selected roll call's complete ordered voting sequence and matching it to the drawn country positions and topic covariates. All parameters are then re-estimated. The TBIP model itself is not re-estimated in each replication.

The reported point estimates remain the original-sample MLEs evaluated at the mean-based TBIP inputs. Bootstrap standard errors are the sample standard deviations of the replicated estimates.

Active social learning and private information. The estimated learning weight is $\widehat{\kappa}_L = 0.05692$, with bootstrap standard error 0.00224 and 95% percentile interval $[0.05236, 0.06093]$. Within the maintained specification, this supports positive updating from predecessors' informational content. With the sign convention in (48), evidence favoring state one lowers L_{it}^{-}; a positive learning weight then lowers both signal cutoffs in (43). The weight is below the unit-weight benchmark, but it is not a percentage-point effect on voting probabilities.

P5 membership, trade power, diplomatic capacity, and being named in the codebook have negative signal-scale coefficients, with intervals excluding zero. These estimates associate those characteristics with more precise private information, conditional on the other covariates. P5 membership is associated with approximately 26% lower signal standard deviation; a one-standard-deviation increase in trade power or diplomatic capacity with approximately 13% or 21% lower signal standard deviation, respectively; and being named in the description with approximately 39% lower signal standard deviation. First Committee, Third Committee, and topic concentration have positive signal-scale coefficients, with intervals excluding zero. The procedural coefficient in the baseline abstention index is negative, whereas the sponsor-count interval includes zero. The Third Committee and topic-concentration coefficients in the prior equation also have intervals including zero.

Heterogeneity across P5 senders. Table 7 shows negative decisive-vote P5 effects for the USA, UK, and France, and positive for the USSR and ROC. All five estimates have statistical significance.

Table 5: Learning, private information, prior, and baseline abstention.

Parameter or covariate	Estimate	Bootstrap SE	95% percentile interval
Learning weight			
κ_L	0.05692	0.00224	$[0.05236, 0.06093]$
Private-signal scale: $\log \sigma_{it}$			
Intercept	1.66984	0.05252	$[1.55266, 1.75986]$
First Committee	0.04490	0.02274	$[0.00321, 0.09125]$
Third Committee	0.06572	0.01821	$[0.02686, 0.09809]$
Topic concentration (standardized)	0.02425	0.00679	$[0.00378, 0.02977]$
P5 member	−0.30708	0.01401	$[-0.33086, -0.27713]$
Trade power (standardized)	−0.13449	0.00710	$[-0.14849, -0.11951]$
Diplomatic capacity (standardized)	−0.23391	0.00727	$[-0.24777, -0.21986]$
Country named in codebook	−0.49288	0.03145	$[-0.55192, -0.43061]$
Prior: $\log[\rho_t/(1-\rho_t)]$			
Intercept	−0.10062	0.05632	$[-0.19628, 0.01890]$
First Committee	−0.04847	0.01071	$[-0.06541, -0.02570]$
Third Committee	−0.00045	0.00824	$[-0.01148, 0.02114]$
Topic concentration (standardized)	−0.01143	0.00331	$[-0.01094, 0.00222]$
Baseline abstention index			
Intercept	−0.00375	0.00145	$[-0.00645, -0.00086]$
Procedural vote	−0.00564	0.00232	$[-0.00998, -0.00137]$
$\log(1 + \text{Sponsors})$ (standardized)	−0.00012	0.00028	$[-0.00070, 0.00041]$

Notes: Uncertainty uses 480 accepted of 500 combined TBIP-posterior-draw and roll-call bootstrap replications. Positive signal-scale coefficients increase private-signal noise. Baseline abstention coefficients enter before the softplus transformation. All preference coefficients are jointly estimated but omitted from this table.

Table 6: Bilateral influence coefficients.

	Decisive-vote channel			Abstention channel		
Feature	Estimate	Boot. SE	95% interval	Estimate	Boot. SE	95% interval
P5 asymmetry: USA reference	−6.0541	0.8914	$[-8.0175, -4.6292]$	−1.4611	1.0261	$[-3.6117, 0.5835]$
Trade-power asymmetry	0.1061	0.0104	$[0.0886, 0.1280]$	0.0765	0.0175	$[0.0404, 0.1104]$
Bilateral trade	0.0650	0.0297	$[0.0063, 0.1196]$	0.3849	0.0400	$[0.3094, 0.4650]$
Diplomatic representation	1.1330	0.0689	$[1.0222, 1.2939]$	0.9433	0.0411	$[0.8789, 1.0344]$
Colonial tie	−0.5515	0.2102	$[-0.9740, -0.1728]$	1.7280	0.3419	$[1.1079, 2.5114]$
USSR deviation from USA	14.0979	1.3142	$[11.8085, 17.0567]$	8.6003	1.5117	$[5.7551, 11.4667]$
UK deviation from USA	−9.0452	1.5582	$[-12.1479, -5.9019]$	3.1979	1.7087	$[-0.1487, 6.5651]$
France deviation from USA	−3.6810	0.9475	$[-5.4626, -1.7660]$	1.5430	1.2299	$[-0.8390, 3.9503]$
ROC deviation from USA	10.5186	1.1497	$[8.5684, 12.9347]$	0.8086	1.2059	$[-1.4247, 3.3856]$

Notes: All numerical entries are multiplied by 10^3. Intervals are the empirical 2.5th and 97.5th percentiles of the combined bootstrap estimates. Continuous bilateral features are scaled without centering; binary features remain unscaled. Decisive weights multiply signed predecessor Yes/No votes; abstention weights multiply predecessor abstentions. These coefficients are not marginal probability effects.

Table 7: P5-to-non-P5 components: reference coefficient plus sender deviation.

	Decisive-vote channel			Abstention channel		
Sender	Estimate	Boot. SE	95% interval	Estimate	Boot. SE	95% interval
USA	−6.0541	0.8914	[−8.0175, −4.6292]	−1.4611	1.0261	[−3.6117, 0.5835]
USSR	8.0438	0.8782	[6.5112, 9.7844]	7.1393	1.0909	[4.9935, 9.3101]
UK	−15.0992	1.2285	[−17.7208, −12.8859]	1.7368	0.9694	[−0.1776, 3.5765]
France	−9.7351	0.7770	[−11.4242, −8.5094]	0.0819	0.7813	[−1.5229, 1.5721]
ROC (China identifier)	4.4645	0.5903	[3.4407, 5.6947]	−0.6525	0.6774	[−1.9926, 0.6412]

Notes: All numerical entries are multiplied by 10^3. Each interval is computed after summing the reference and deviation coefficients within each bootstrap replication, retaining their covariance. These are the P5-specific components $b_c + d_{c,s}$, not the complete bilateral weights $\alpha^c_{il,t}$.

The trade-power asymmetry, bilateral trade, and diplomatic representation effects for decisive vote channel are positive and colonial ties have negative effects. These components must be combined with the P5 terms to describe any particular country pair effect. The estimates describe conditional influence within the model, rather than establishing a general causal tendency to oppose a particular state.

Abstention and political cover. Trade-power asymmetry, bilateral trade, diplomatic representation, and colonial ties have positive abstention-channel coefficients with statistical significance. Among the five P5 components, only the USSR has statistical significance, with estimate being 0.007139, with interval $[0.004993, 0.009310]$.

The results therefore support relationship-specific political cover, with more limited evidence for a common P5 abstention effect. Bilateral weights enter before the softplus transformation and accumulate over predecessors; their magnitudes cannot be read as probability changes or compared directly with the learning coefficient. Positive social learning remains present after allowing both political-influence channels to vary across P5 senders.

Table 8 summarizes, for each P5 sender, the number of recipient countries with positive, negative, or zero exposure-weighted average bilateral influence in each channel, combining all relationship features and sender deviations.

Table 8: Signs of complete P5 bilateral weights at the fitted MLE.

	Decisive-vote channel			Abstention channel		
Sender	Positive	Negative	Zero	Positive	Negative	Zero
USA	4	122	0	113	13	0
USSR	126	0	0	126	0	0
UK	4	122	0	126	0	0
France	4	121	0	125	0	0
ROC	125	0	1	18	107	1

Notes: Each entry counts recipient voting identifiers, excluding self-links and including only pairs observed with the sender preceding the recipient. For each pair, year-specific complete weights are averaged using the number of observed predecessor exposures as weights. Signs refer to these averages of $\alpha^{\text{thr}}_{il,t}$ and $\alpha^{d}_{il,t}$, including all bilateral features and sender deviations, not just the P5 components. They are neither counts of statistically significant links nor signs of realized vote-weighted contributions. Zero uses tolerance 10^{-12}. Egypt and UAR retain separate voting identifiers.

Appendix C reports Bayesian estimates alongside the MLE results as an estimation cross-check under the maintained specification.

Appendix D considers alternatives using country intercepts and codebook-derived covariates, with and without country-specific committee slopes. These specifications require no TBIP first stage and hence no first-stage uncertainty correction, although sampling uncertainty remains.

Appendix E uses ideal points estimated from General Debate speeches rather than resolution-specific speech. Together, these alternatives assess sensitivity to reliance on potentially strategic roll-call rhetoric; they do not eliminate strategic communication, which may also shape General Debate statements.

Counterfactual Voting Outcomes

Appendix F presents exploratory counterfactual exercises illustrating the implications of the current specification. These preliminary results condition on the fitted parameters; their sensitivity to alternative specifications and estimation uncertainty remains to be assessed.

Conclusion

I study how sequential voting shapes behavior in collective choice. When votes are publicly observed in sequential voting, earlier vote decisions may affect later voters by revealing information about the underlying proposal or exerting pressure via voters' relationships outside the voting process. Identifying sequential influence and distinguishing the underlying mechanisms are difficult because voting order is often endogenous to institutional rules or strategic behavior.

I examine this question using United Nations General Assembly roll calls during the early Cold War. For reduced-form, I use country- and roll-call-fixed-effects regressions to test whether countries align more closely with particular states that vote earlier. The results show heterogeneous effects: countries are more likely to align with global or regional powers, and the effects vary across those countries.

Additional tests suggest positive influence through both learning and relationships outside the voting process, with stronger support for the latter.

I further estimate a structural model. Conditional on the model assumptions, the estimates indicate positive informational updating and heterogeneous bilateral influence from outside the voting process. Exploratory counterfactual simulations compare sequential and simultaneous voting and remove informational or bilateral influence to assess changes in vote shares and decision-threshold outcomes.

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A Derivation of the signal thresholds

This appendix derives the signal-threshold representation (43) from the posterior thresholds (34), and establishes the width and asymmetry properties stated in the text.

Posterior log-odds. Country i's information set Γ_{it} comprises the prior on ω_t, the public votes of the predecessors $\mathcal{A}_{it}$, and its own private signal s_{it}. Taking the private signal to be conditionally independent of the public history given ω_t, the posterior log-odds decomposes additively into a prior, a public, and a private term:

$$\log \frac{q_{it}}{1-q_{it}} = \log \frac{\rho_t}{1-\rho_t} + \ell_{it}^{\text{pub}} + \ell_{it}^{\text{priv}}, \tag{58}$$

where each likelihood-ratio term is written for $\omega_t = 1$ against $\omega_t = 0$.

The private signal. Under $s_{it} \mid \omega_t \sim \mathcal{N}(\omega_t, \sigma_{it}^2)$, the density is $f(s \mid \omega_t) = (2\pi\sigma_{it}^2)^{-1/2} \exp\{-(s-\omega_t)^2/2\sigma_{it}^2\}$, so

$$\ell_{it}^{\text{priv}} = \log \frac{f(s_{it} \mid \omega_t = 1)}{f(s_{it} \mid \omega_t = 0)} = \frac{1}{2\sigma_{it}^2}\left[\, s_{it}^2 - (s_{it}-1)^2 \,\right] = \frac{s_{it} - \frac{1}{2}}{\sigma_{it}^2}. \tag{59}$$

The signal enters the log-odds linearly, with slope $1/\sigma_{it}^2$: a sharper signal (smaller σ_{it}^2) moves the posterior more per unit of s_{it}.

The public history. Let

$$L_{it}^{-} = \sum_{l \in \mathcal{A}_{it}} \log \frac{\Pr(v_{lt} \mid \omega_t = 0)}{\Pr(v_{lt} \mid \omega_t = 1)} \tag{60}$$

denote the accumulated public log-likelihood ratio in favour of $\omega_t = 0$, with κ_L the weight placed on this social information. Because (60) is written for state 0, its contribution to the log-odds for state 1 carries the opposite sign, $\ell_{it}^{\text{pub}} = -\kappa_L L_{it}^{-}$.

Inversion. Substituting (59) and $\ell_{it}^{\text{pub}} = -\kappa_L L_{it}^{-}$ into (58),

$$\log \frac{q_{it}}{1-q_{it}} = \log \frac{\rho_t}{1-\rho_t} - \kappa_L L_{it}^{-} + \frac{s_{it} - \frac{1}{2}}{\sigma_{it}^2}. \tag{61}$$

Fix a threshold $v \in \{N, Y\}$ and let s_{it}^v be the signal at which the posterior equals the corresponding cutoff, $q_{it} = q_{it}^v$. Solving (61) for s_{it}^v,

$$s_{it}^v = \frac{1}{2} + \sigma_{it}^2 \left[\log \frac{q_{it}^v}{1-q_{it}^v} - \log \frac{\rho_t}{1-\rho_t} + \kappa_L L_{it}^{-} \right] = \frac{1}{2} + \sigma_{it}^2 \left[\log \frac{q_{it}^v}{1-q_{it}^v} + \log \frac{1-\rho_t}{\rho_t} + \kappa_L L_{it}^{-} \right], \tag{62}$$

which is (43). Since the map $s \mapsto q$ implied by (61) is strictly increasing, the ordering $q_{it}^N \le q_{it}^Y$ carries over to $s_{it}^N \le s_{it}^Y$, and the posterior rule (35) is equivalent to the signal rule (44).

B Monte Carlo evidence

I validate estimation with three exercises. The first is an identification Monte Carlo on a stylized, few-type design, which confirms that the structural parameters are separately identified. The second is a maximum-likelihood recovery study on the full specification (31 parameters, $N = 100$ countries, $T = 3000$ roll calls), which checks the estimator that produces the point estimates. The third is a simulation-based calibration, which checks that the Bayesian sampler returns a correctly calibrated posterior.

B.1 Identification

I first confirm that the structural parameters are separately identified in a deliberately stylized version of the model. There are $N = 9$ countries, divided evenly among three discrete types, A, B, and C. Country type determines the signal noise σ_i and preference bias π_i. Pairwise relations are reduced to three types: same-type pairs, A–C pairs, and a baseline consisting of A–B and B–C pairs. Voting order is independently randomized across roll calls. Each country observes a noisy private signal about a binary resolution state and then chooses No, Abstain, or Yes. Earlier decisive votes shift the Yes–No threshold through the pair-specific coefficients α_{thr}, earlier abstentions shift the width of the abstention region through α_{band}, and the informational content of preceding votes enters through the accumulated likelihood ratio with weight κ_L. The parameter τ determines the baseline width of the abstention region. Thus the exercise retains the main preference, precision, learning, decisive-pressure, and abstention channels of the full model while making their sources of variation transparent.

Across $R = 1000$ simulated panels of $T = 300$ roll calls, I estimate the model from five different starting values and retain the best-converged optimum. The estimation uses diffuse priors so that the resulting MAP estimates closely approximate maximum likelihood; standard errors are computed from the Hessian of the likelihood. Figure 5 plots the resulting sampling distributions. The three signal precisions and three preference biases separate cleanly; the positive and negative coefficients in both the decisive-pressure and abstention channels recover their respective truths; and the learning weight and baseline abstention width are centered on theirs. The distinct channels are therefore recovered without substantial tradeoffs, providing a finite-sample counterpart to the identification argument in the Identification Section.

B.2 Maximum-likelihood recovery

The full specification has 31 parameters. Four quantities vary with covariates through log- or logit-linear links, each built from a roll-call block $\mathbf{x}_t$, a country block $\mathbf{z}_i$, and a country–roll-call block $\mathbf{w}_{it}$:

$$\log \sigma_{it} = \mathbf{x}_t^\top \boldsymbol{\alpha}_\sigma + \mathbf{z}_i^\top \boldsymbol{\beta}_\sigma + \mathbf{w}_{it}^\top \chi_\sigma, \qquad 1 - \pi_{it} = \tfrac{1}{2} + \mathbf{x}_t^\top \boldsymbol{\zeta}_\pi + \mathbf{z}_i^\top \boldsymbol{\eta}_\pi + \mathbf{w}_{it}^\top \lambda_\pi,$$
$$\theta_{d,it} = \mathbf{x}_t^\top \boldsymbol{\theta}_X + \mathbf{z}_i^\top \boldsymbol{\theta}_Z + \mathbf{w}_{it}^\top \theta_W, \qquad \log \frac{\rho_t}{1-\rho_t} = \mathbf{x}_t^\top \boldsymbol{\delta}_\rho.$$

The bias link generalizes the spatial form of $1 - \pi_{it}$ in the model section; the two alliance coefficients are linear in directed pair features $\mathbf{g}_{il}$, with $\alpha_{il}^{\text{thr}} = \mathbf{g}_{il}^\top \boldsymbol{\beta}_{\text{thr}}$ and $\alpha_{il}^{d} = \mathbf{g}_{il}^\top \boldsymbol{\beta}_d$. Together with the learning weight κ_L, the estimated vector is $\theta = (\boldsymbol{\alpha}_\sigma, \boldsymbol{\beta}_\sigma, \chi_\sigma, \boldsymbol{\zeta}_\pi, \boldsymbol{\eta}_\pi, \lambda_\pi, \boldsymbol{\delta}_\rho, \boldsymbol{\theta}_X, \boldsymbol{\theta}_Z, \theta_W, \boldsymbol{\beta}_{\text{thr}}, \boldsymbol{\beta}_d, \kappa_L)$, whose blocks label the panels of Figure 6 and the rows of Table 9.

For each replication I draw a panel at the true θ, estimate θ by maximum likelihood from multiple random starts (retaining the best-converged optimum), and form Wald confidence intervals from the inverse observed information. Figure 6 plots the sampling distribution of each estimate, over the $R = 967$ converged replications, against its true value. Three features stand out. Bias is negligible for every parameter: the mean estimate and the true value coincide to plotting accuracy. Ninety-five percent Wald coverage is close to nominal, between 0.94 and 0.96 across parameters. And the ratio of the mean estimated standard error to the empirical standard deviation of the estimates is approximately one, so the information-based standard errors are reliable. The low-frequency alliance and wedge coefficients $\boldsymbol{\beta}_{\text{thr}}$ and $\boldsymbol{\beta}_d$, identified only by the relatively rare histories in which an ally has already committed or abstained, are recovered without bias at this sample size.

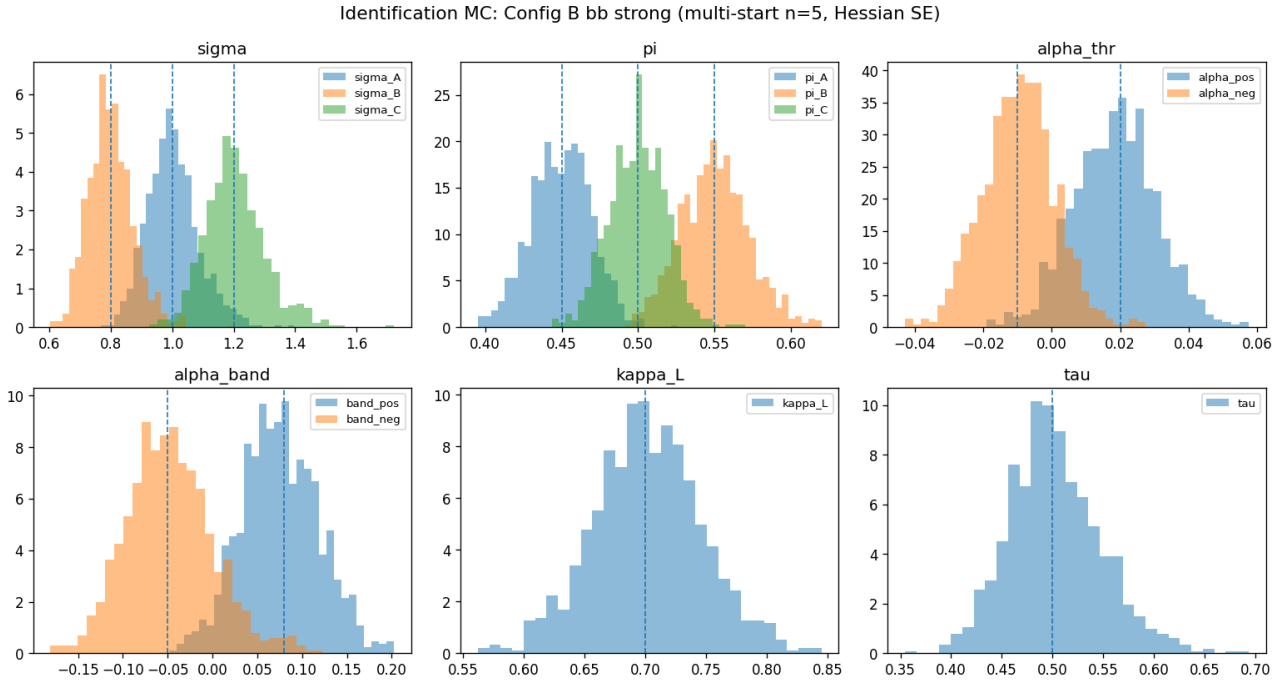


Figure 5: Identification Monte Carlo in the stylized three-type model. There are $N = 9$ countries (three each of types A, B, and C) and $T = 300$ roll calls per replication, with voting order randomly permuted across roll calls. The panels report the sampling distributions across $R = 1000$ replications for the type-specific signal noise σ, type-specific preference bias π, the positive- and negative-pair coefficients in the decisive-pressure channel α_{thr}, the corresponding coefficients in the abstention channel α_{band}, the learning weight κ_L, and the baseline abstention width τ. Each replication is estimated from five starting values, retaining the best-converged optimum; Hessian standard errors use the curvature of the likelihood. Dashed vertical lines indicate the true parameter values. The sampling distributions are centered closely on their respective truths.

B.3 Simulation-Based Calibration

The Bayesian analysis requires a distinct check: that the posterior sampler returns the correct posterior, not merely a point estimate near the truth. I use the simulation-based calibration (SBC) rank test of Talts et al. (2020), which rests on the fact that a parameter drawn from the prior, passed through the likelihood and then the posterior, should be statistically indistinguishable from the posterior draws that surround it. For each replication I (i) draw $\tilde{\theta} \sim \pi(\theta)$; (ii) simulate data $\tilde{y} \sim \pi(y \mid \tilde{\theta})$; (iii) draw $\theta_1, \ldots, \theta_L \sim \pi(\theta \mid \tilde{y})$; and (iv) record the rank of $\tilde{\theta}$ among the L posterior draws. Under exact inference these ranks are uniform on $\{0, \ldots, L\}$; departures from uniformity diagnose a miscalibrated sampler.

I run $R = 1000$ replications with $L = 2000$ posterior draws each; 870 (87%) reach the convergence threshold and enter the analysis. Figure 7 shows the rank histograms, which lie within the uniform band for all 31 parameters, whose blocks are defined at the start of the recovery study above (Appendix B.2). Table 9 summarizes calibration by the standardized deviation of each parameter's mean rank from its uniform expectation (z) and by the empirical 90% credible-interval coverage. Every mean rank sits at 0.5 to within sampling error: the largest deviation is $|z| = 2.16$, for $\beta_\sigma[1]$, which is unremarkable across 31 tests, and 90% coverage lies between 0.88 and 0.93 throughout. Posterior contraction relative to the prior is high for the precision, prior, and learning blocks and low for the weakly identified preference and alliance blocks ($\boldsymbol{\zeta}_\pi$, $\boldsymbol{\beta}_{\text{thr}}$, $\boldsymbol{\beta}_d$), where the identifying histories are rare; calibration holds in both regimes,

Table 9: Simulation-based calibration ($R = 870$ converged replications, ranks among $L = 2000$ posterior draws). z is the standardized deviation of the mean rank from the uniform expectation; cov_{90} is the empirical 90% credible-interval coverage.

Parameter	z	cov_{90}	Parameter	z	cov_{90}
$\alpha_\sigma[0]$	$+0.68$	0.889	$\eta_\pi[0]$	$+0.23$	0.884
$\alpha_\sigma[1]$	$+0.04$	0.903	$\eta_\pi[1]$	-0.53	0.903
$\alpha_\sigma[2]$	$+0.29$	0.898	$\eta_\pi[2]$	-0.32	0.893
$\alpha_\sigma[3]$	$+0.26$	0.910	$\lambda_\pi[0]$	-0.37	0.889
$\beta_\sigma[0]$	-0.40	0.902	$\beta_{\text{thr}}[0]$	-1.00	0.898
$\beta_\sigma[1]$	-2.16	0.900	$\beta_{\text{thr}}[1]$	-0.63	0.901
$\beta_\sigma[2]$	$+1.70$	0.887	κ_L	$+0.53$	0.883
$\chi_\sigma[0]$	-0.65	0.890	$\theta_X[0]$	-1.28	0.909
$\delta_\rho[0]$	-0.51	0.879	$\theta_X[1]$	$+1.90$	0.895
$\delta_\rho[1]$	-0.06	0.897	$\theta_X[2]$	-1.67	0.885
$\delta_\rho[2]$	-0.36	0.894	$\theta_Z[0]$	-0.66	0.897
$\delta_\rho[3]$	-0.17	0.875	$\theta_Z[1]$	-0.53	0.931
$\zeta_\pi[0]$	$+1.54$	0.892	$\theta_W[0]$	-0.48	0.898
$\zeta_\pi[1]$	$+0.31$	0.895	$\beta_d[0]$	-0.53	0.895
$\zeta_\pi[2]$	-0.76	0.887	$\beta_d[1]$	$+0.12$	0.907
$\zeta_\pi[3]$	$+0.36$	0.916			

which is precisely what SBC verifies.

C MLE and Bayesian Estimates of the P5-Sender Specification

The Bayesian run estimates the same 109-parameter ROC/P5-sender model, conditional on the fixed mean-based TBIP covariates. It uses four chains, each with 1,000 warmup iterations and 1,000 retained draws, initialized from the original unfitted MLE starting vector. Priors are specified in optimizer coordinates with support restricted to the MLE parameter bounds. MLE intervals instead use the combined text-posterior/roll-call bootstrap replications, so the two sets of intervals represent different uncertainty calculations. Only coefficients and derived sender components reported in the main estimation section are included; the 75 preference coefficients are omitted.

Table 10: Learning, information, prior, and abstention: MLE and Bayesian estimates.

Parameter	MLE	Bootstrap 95% CI	Posterior mean	95% credible interval
Learning weight κ_L	0.05692	$[0.05236, 0.06093]$	0.05607	$[0.05411, 0.05811]$
Private-signal scale				
Intercept	1.66984	$[1.55266, 1.75986]$	1.58945	$[1.53507, 1.64597]$
First Committee	0.04490	$[0.00321, 0.09125]$	0.04183	$[0.01973, 0.06378]$
Third Committee	0.06572	$[0.02686, 0.09809]$	0.06580	$[0.04514, 0.08724]$
Topic concentration	0.02425	$[0.00378, 0.02977]$	0.02310	$[0.01591, 0.03039]$
P5 membership	−0.30708	$[-0.33086, -0.27713]$	−0.30625	$[-0.33571, -0.27838]$
Trade power	−0.13449	$[-0.14849, -0.11951]$	−0.13596	$[-0.14551, -0.12611]$
Diplomatic capacity	−0.23391	$[-0.24777, -0.21986]$	−0.23219	$[-0.24091, -0.22337]$
Named in description	−0.49288	$[-0.55192, -0.43061]$	−0.48403	$[-0.54531, -0.42199]$
Prior log odds				
Intercept	−0.10062	$[-0.19628, 0.01890]$	−0.06024	$[-0.14110, 0.01906]$
First Committee	−0.04847	$[-0.06541, -0.02570]$	−0.05302	$[-0.06201, -0.04401]$
Third Committee	−0.00045	$[-0.01148, 0.02114]$	−0.00053	$[-0.00889, 0.00791]$
Topic concentration	−0.01143	$[-0.01094, 0.00222]$	−0.01240	$[-0.01593, -0.00888]$
Baseline abstention index				
Intercept	−0.00375	$[-0.00645, -0.00086]$	−0.00160	$[-0.00316, -0.00007]$
Procedural vote	−0.00564	$[-0.00998, -0.00137]$	−0.00582	$[-0.00823, -0.00353]$
Sponsor count (log, standardized)	−0.00012	$[-0.00070, 0.00041]$	−0.00013	$[-0.00043, 0.00016]$

Notes: MLE intervals use 480 accepted original-start combined-bootstrap replications. Bayesian intervals condition on fixed TBIP inputs. Sender totals are summed within each replication or posterior draw before taking quantiles.

Table 11: Decisive-vote bilateral coefficients: MLE and Bayesian estimates.

Parameter	MLE	Bootstrap 95% CI	Posterior mean	95% credible interval
USA reference	−6.0541	$[-8.0175, -4.6292]$	−6.5539	$[-7.8745, -5.2319]$
Trade-power asymmetry	0.1061	$[0.0886, 0.1280]$	0.1129	$[0.0969, 0.1294]$
Bilateral trade	0.0650	$[0.0063, 0.1196]$	0.0628	$[0.0303, 0.0959]$
Diplomatic representation	1.1330	$[1.0222, 1.2939]$	1.2291	$[1.1516, 1.3071]$
Colonial tie	−0.5515	$[-0.9740, -0.1728]$	−0.5851	$[-0.9092, -0.2826]$
USSR minus USA	14.0979	$[11.8085, 17.0567]$	15.3528	$[13.6256, 17.1210]$
UK minus USA	−9.0452	$[-12.1479, -5.9019]$	−9.9183	$[-12.2413, -7.5383]$
France minus USA	−3.6810	$[-5.4626, -1.7660]$	−4.1174	$[-5.6117, -2.5803]$
ROC minus USA	10.5186	$[8.5684, 12.9347]$	11.3461	$[9.6650, 13.0422]$

Notes: All numerical entries are multiplied by 10^3. MLE intervals use 480 accepted original-start combined-bootstrap replications. Bayesian intervals condition on fixed TBIP inputs. Sender totals are summed within each replication or posterior draw before taking quantiles.

Table 12: Abstention bilateral coefficients: MLE and Bayesian estimates.

Parameter	MLE	Bootstrap 95% CI	Posterior mean	95% credible interval
USA reference	−1.4611	[−3.6117, 0.5835]	−1.5435	[−2.7565, −0.3008]
Trade-power asymmetry	0.0765	[0.0404, 0.1104]	0.0795	[0.0551, 0.1042]
Bilateral trade	0.3849	[0.3094, 0.4650]	0.4061	[0.3578, 0.4555]
Diplomatic representation	0.9433	[0.8789, 1.0344]	0.9848	[0.9300, 1.0418]
Colonial tie	1.7280	[1.1079, 2.5114]	1.8822	[1.3191, 2.4454]
USSR minus USA	8.6003	[5.7551, 11.4667]	8.9622	[7.1739, 10.7452]
UK minus USA	3.1979	[−0.1487, 6.5651]	3.2191	[1.1828, 5.2068]
France minus USA	1.5430	[−0.8390, 3.9503]	1.5499	[−0.0320, 3.0857]
ROC minus USA	0.8086	[−1.4247, 3.3856]	0.8169	[−0.7822, 2.3095]

Notes: All numerical entries are multiplied by 10^3. MLE intervals use 480 accepted original-start combined-bootstrap replications. Bayesian intervals condition on fixed TBIP inputs. Sender totals are summed within each replication or posterior draw before taking quantiles.

Table 13: P5-to-non-P5 components: reference plus sender deviation.

Parameter	MLE	Bootstrap 95% CI	Posterior mean	95% credible interval
Decisive-vote channel				
USA	−6.0541	[−8.0175, −4.6292]	−6.5539	[−7.8745, −5.2319]
USSR	8.0438	[6.5112, 9.7844]	8.7990	[7.7062, 9.9667]
UK	−15.0992	[−17.7208, −12.8859]	−16.4722	[−18.1362, −14.7715]
France	−9.7351	[−11.4242, −8.5094]	−10.6713	[−11.7820, −9.5616]
ROC	4.4645	[3.4407, 5.6947]	4.7923	[3.8872, 5.7189]
Abstention channel				
USA	−1.4611	[−3.6117, 0.5835]	−1.5435	[−2.7565, −0.3008]
USSR	7.1393	[4.9935, 9.3101]	7.4187	[6.1242, 8.6975]
UK	1.7368	[−0.1776, 3.5765]	1.6756	[0.4567, 2.8795]
France	0.0819	[−1.5229, 1.5721]	0.0064	[−0.9337, 0.9374]
ROC	−0.6525	[−1.9926, 0.6412]	−0.7265	[−1.5936, 0.1223]

Notes: All numerical entries are multiplied by 10^3. MLE intervals use 480 accepted original-start combined-bootstrap replications. Bayesian intervals condition on fixed TBIP inputs. Sender totals are summed within each replication or posterior draw before taking quantiles.

D Country Intercepts and Committee-Specific Preference Slopes

As an exploratory preference specification, I replace the TBIP ideal-point and topic-share terms in the preference equation with country intercepts, codebook-derived content indicators, and country-specific committee slopes:

$$\text{logit}(\pi_{it}) = a_i + \mathbf{C}_t^\top \boldsymbol{\zeta} + \chi \, Named_{it} + b_{i1} C1_t + b_{i3} C3_t.$$

Here $\mathbf{C}_t$ contains 12 indicators for issue content, proposed actions, ambiguous action wording, and missing descriptions; $Named_{it}$ indicates that country i is named in the description. The interactions are between country indicators and First/Third Committee status, rather than between country indicators and every codebook variable. Other committee assignments form the reference category. The codebook measures are retrospective descriptions, not verified pre-vote information.

The sample remains 206,625 decisions in 2,679 roll calls over 1948–1971. This alternative retains the corrected ROC bilateral mapping and uses directional recipient trade dependence, P5 sender effects, and defense-pact heterogeneity in both influence channels. It therefore also differs from the main specification in its bilateral covariates. In both alternatives, TBIP topic concentration is removed from the signal-scale and prior equations; their resolution-level covariates are an intercept and First/Third Committee indicators. No TBIP covariates enter any equation in these alternatives. The common estimation sample nevertheless retains the earlier text-matching restriction. Adding 254 country–committee slopes to the 182-parameter country-intercept/codebook model yields 436 parameters.

Comparison with the TBIP preference specification. Table 14 compares the main 109-parameter TBIP model with two alternatives: country intercepts and common codebook slopes (182 parameters), and the same specification augmented with country–committee slopes (436 parameters). Both alternatives use the same directional-trade and defense-pact variables. Comparing the last two columns therefore isolates the addition of committee slopes; comparing either with the TBIP column combines preference and bilateral-specification changes with the removal of signal/prior topic concentration. These are point-estimate comparisons, not tests of equality across specifications.

The learning estimate remains positive, increasing from 0.05692 in the TBIP specification to 0.08529 with country/codebook preferences and 0.08586 after adding committee slopes. Thus the larger change occurs before introducing the committee interactions. The negative signal-scale coefficients on P5 membership, trade power, diplomatic capacity, and being named also persist. Their magnitudes change, however: the trade-power and diplomatic-capacity coefficients are substantially closer to zero in the alternatives. The signs are consistent with the same qualitative information ranking within each model, without establishing statistical robustness across specifications.

The committee coefficients illustrate a less stable decomposition. In the TBIP model the First and Third Committee signal-scale coefficients are positive; with country–committee preference slopes they become negative. The committee coefficients in the prior equation also become more negative. Allowing country-specific preference shifts therefore changes how the model allocates committee-related variation across preferences, prior beliefs, and signal precision. These movements should not be read as independently identified changes in information quality. The procedural and sponsor-count point estimates in the baseline abstention equation remain negative.

Table 14: Key structural estimates across preference specifications.

Parameters	TBIP 109	Country/codebook 182	Country/codebook + committee slopes 436
Learning weight κ_L	0.05692	0.08529	0.08586
Private-signal scale: $\log \sigma_{it}$			
P5 membership	−0.30708	−0.28216	−0.28676
Trade power	−0.13449	−0.02614	−0.02612
Diplomatic capacity	−0.23391	−0.07814	−0.07876
Named in description	−0.49288	−0.56367	−0.55245
First Committee	0.04490	0.05825	−0.01184
Third Committee	0.06572	0.02588	−0.05490
Topic concentration	0.02425	–	–
Prior log odds			
First Committee	−0.04847	−0.04681	−0.32193
Third Committee	−0.00045	−0.00141	−0.33390
Baseline abstention index			
Intercept	−0.00375	−0.00331	−0.00331
Procedural vote	−0.00564	−0.00697	−0.00699
Sponsor count (log, standardized)	−0.00012	−0.00039	−0.00039

Notes: All models use the same voting sample and corrected ROC mapping. TBIP denotes the main P5-sender specification based on roll-call text, not General Debate ideal points. Signal/prior topic concentration is retained only in the main TBIP model; dashes indicate exclusion. Only point estimates are shown: the 436-parameter fit does not provide usable Hessian-based standard errors.

Bilateral influence. Table 15 reports selected complete P5-asymmetry components, formed by adding the USA reference coefficient to each sender deviation, together with other bilateral coefficients. The sender components apply to non-P5 recipients and exclude trade, diplomacy, colonial, and defense-pact contributions. They are not total country-pair effects.

The USA, UK, and France retain negative decisive-vote P5 components, while the USSR component remains positive. The ROC component changes from positive in the TBIP specification to slightly negative in the alternatives. Diplomatic representation remains positive, but its coefficient is appreciably smaller; the decisive colonial-tie coefficient changes from negative to a small positive value. Most bilateral point estimates change little between the two codebook alternatives. In the abstention channel, the USSR component stays positive, whereas the UK and France components move close to or below zero. These results show that positive learning and several information-coefficient signs persist, while quantitative bilateral effects and some sender-component signs remain sensitive to specification. They do not imply that every recipient faces a bilateral weight of the same sign as its sender component.

The best log-likelihood is −145,971.82, compared with −146,350.97 without country–committee slopes. The improvement is an in-sample fit comparison and does not establish better predictive performance. Selected P5 committee slopes are reported below.

All six optimizer starts satisfy the relative objective-reduction stopping rule and produce log-likelihoods within 0.0022 of one another. The scaled projected-gradient norm is approximately 0.00767, and no coefficient is at an optimizer bound. Nevertheless, the observed information matrix is numerically singular (minimum eigenvalue approximately -4.3×10^{-12}) and fails the positive-definiteness check. Agreement across starts does not resolve this flat curvature. Hessian standard errors and confidence intervals are unavailable for the 436-parameter fit. The 182-parameter fit has a positive-definite information matrix, no active bounds, and a scaled projected-gradient norm of approximately 0.00536. The committee extension is therefore reported as an exploratory point-estimate comparison; the source of its weak curvature requires

Table 15: Selected bilateral estimates across specifications ($\times 10^3$).

	TBIP	Country/codebook	Country/codebook + committee slopes
Decisive-vote channel			
USA component	−6.054	−11.606	−11.640
USSR component	8.044	3.287	3.436
UK component	−15.099	−12.456	−12.461
France component	−9.735	−11.864	−11.949
ROC component	4.465	−0.162	−0.208
Trade-power asymmetry	0.106	0.150	0.149
Diplomatic representation	1.133	0.205	0.207
Colonial tie	−0.552	0.074	0.069
Log bilateral trade	0.065	–	–
Recipient trade dependence	–	0.874	0.869
Defense pact (common component)	–	5.296	5.299
Abstention channel			
USA component	−1.461	−7.253	−7.207
USSR component	7.139	3.163	3.158
UK component	1.737	−0.167	−0.206
France component	0.082	−0.827	−0.789
ROC component	−0.653	−0.492	−0.481
Trade-power asymmetry	0.076	0.203	0.201
Diplomatic representation	0.943	0.392	0.391
Colonial tie	1.728	0.593	0.646
Log bilateral trade	0.385	–	–
Recipient trade dependence	–	0.971	0.958
Defense pact (common component)	–	4.682	4.683

Notes: All estimates are multiplied by 10^3. A dash denotes an excluded regressor. Log trade and recipient trade dependence are different variables with different scales; their magnitudes are not directly comparable. The alternatives also include sender-specific P5–defense-pact interactions, omitted here for brevity. Entries are index coefficients, not marginal effects on voting probabilities.

Table 16: Selected country-specific committee slopes in the preference index.

Country	First Committee	Third Committee
USA	0.23175	0.26927
USSR	0.32437	0.38757
UK	0.26301	0.34997
France	0.30828	0.41355
ROC (China identifier)	0.23986	0.33340

Notes: Entries are coefficients in $\text{logit}(\pi_{it})$ relative to other committee assignments, conditional on country intercepts and codebook content. They are not changes in Yes-vote probabilities. Only point estimates are reported because the fitted information matrix is numerically singular.

further diagnosis before using it for inference.

Signs of complete bilateral links. Table 17 counts recipient countries with positive, negative, or zero complete bilateral weights for each P5 sender. Each link combines all relationship features, sender deviations, and defense-pact interactions and is averaged across years using observed predecessor exposures. These are signs of fitted weights, not tests of statistical significance or counts of realized voting responses.

Table 17: Signs of complete P5 bilateral links under the revised codebook specifications.

	Decisive-vote channel			Abstention channel		
Sender	Positive	Negative	Zero	Positive	Negative	Zero
Country/codebook (182 parameters)						
USA	42	84	0	44	82	0
USSR	126	0	0	126	0	0
UK	25	101	0	126	0	0
France	29	96	0	122	3	0
ROC	56	69	1	23	102	1
Country/codebook + committee slopes (436 parameters)						
USA	42	84	0	44	82	0
USSR	126	0	0	126	0	0
UK	25	101	0	126	0	0
France	29	96	0	122	3	0
ROC	46	79	1	23	102	1

Notes: Counts exclude self-links and include only observed sender–recipient pairs, including P5 recipients. Absolute weights at most 10^{-12} are classified as zero. Egypt and the United Arab Republic retain distinct voting identifiers. Positive decisive weights favor alignment with a predecessor's Yes/No position; positive abstention weights increase the abstention incentive following that sender's abstention. Counts reflect point estimates and do not incorporate estimation uncertainty.

Adding committee slopes leaves all abstention counts and all non-ROC decisive counts unchanged. For ROC, ten recipient links change from positive to negative: positive links fall from 56 to 46 and negative links rise from 69 to 79. A sender's P5 component alone therefore does not determine the sign of its complete bilateral links.

E General Debate Ideal Points and Codebook Content

This appendix considers an alternative source of text-based country positions. The United Nations General Debate corpus contains national statements delivered in the General Assembly's annual General Debate, covering governments' broad international priorities and policy positions. These statements are distinct from speeches attached to particular roll-call votes. I use the 1948–1971 portion of the available corpus and fit a two-dimensional TBIP model with 25 latent topics by variational inference. Country coordinates are pooled across the period rather than estimated separately for each year.

Using General Debate speeches reduces reliance on resolution-specific rhetoric that may reflect the voting process being modeled. It does not make the text nonstrategic: national statements can themselves serve signaling and diplomatic purposes. Moreover, pooled coordinates use later speeches to measure positions for earlier votes and are not predetermined at each vote date. This is a measurement sensitivity exercise, not a solution to all concerns about strategic communication or endogeneity.

Specification. Let $\widehat{\mathbf{x}}_i^{GD}$ denote the two fitted country coordinates and let $\mathbf{C}_t$ contain the 12 codebook-derived issue, action, ambiguity, and missing-description indicators. Preferences satisfy

$$\text{logit}(\pi_{it}) = a_0 + (\widehat{\mathbf{x}}_i^{GD})^\top \boldsymbol{\eta} + \mathbf{C}_t^\top \boldsymbol{\zeta} + \chi Named_{it} + (\widehat{\mathbf{x}}_i^{GD})^\top \Lambda \mathbf{C}_t.$$

There are no country intercepts in this specification. GD topic shares are not used as roll-call content measures; the codebook indicators provide that variation. No covariates from the original roll-call-text TBIP fit enter any equation, and topic concentration is excluded from both signal scale and the prior. Their resolution-level covariates are an intercept and First/Third Committee indicators. The remaining country-level information covariates and the baseline abstention specification are retained. The codebook descriptions are retrospective content proxies, not verified pre-vote documents.

The model retains directional recipient trade dependence, P5 sender heterogeneity, and defense-pact interactions, with the corrected ROC bilateral mapping. There are 82 structural parameters and the same 206,625 decisions in 2,679 roll calls. The inherited sample still reflects the earlier text-matching restriction. Relative to the main 109-parameter model, this comparison changes preference measurement, bilateral covariates, and signal/prior controls together; it does not isolate the effect of changing the text corpus alone.

Estimates and uncertainty. Point estimates use the fitted GD posterior means. For uncertainty, 500 combined replications draw GD country coordinates from the fitted variational posterior, preserving the coordinate orientation and shared country aliases, and resample complete ordered roll calls with replacement. All structural parameters are re-estimated starting from the original-sample fitted MLE. No original roll-call TBIP quantities are drawn, and the GD model is not re-estimated. All 500 replications passed the numerical acceptance criteria. Intervals are the empirical 2.5th and 97.5th percentiles of the accepted estimates; this procedure does not account for variational approximation error.

Table 18 compares the two specifications. Positive learning and negative signal-scale coefficients on P5 membership, trade power, diplomatic capacity, and named-country status persist in this specification. Their bootstrap intervals exclude zero. These patterns provide a sensitivity comparison rather than independent validation of the structural interpretation.

All six optimizer starts converge to closely agreeing objective values, and the observed information matrix is positive definite. However, the USSR–defense-pact abstention interaction reaches its upper bound of 0.05 in the original fit and in 344 of 500 bootstrap replications

Table 18: Selected estimates: main roll-call TBIP and General Debate specifications.

	Main TBIP (109 parameters)		GD/codebook (82 parameters)	
Parameter	MLE	95% bootstrap interval	MLE	95% bootstrap interval
Learning weight κ_L	0.056923	$[0.052357, 0.060926]$	0.081768	$[0.078426, 0.085136]$
Signal: P5 membership	−0.307080	$[-0.330861, -0.277128]$	−0.255761	$[-0.289389, -0.219949]$
Signal: trade power	−0.134492	$[-0.148488, -0.119512]$	−0.024621	$[-0.035133, -0.013152]$
Signal: diplomatic capacity	−0.233912	$[-0.247766, -0.219861]$	−0.106312	$[-0.120500, -0.094024]$
Signal: named in description	−0.492884	$[-0.551921, -0.430611]$	−0.621729	$[-0.694533, -0.553787]$
Signal: First Committee	0.044900	$[0.003215, 0.091249]$	0.070923	$[0.033991, 0.107747]$
Signal: Third Committee	0.065720	$[0.026863, 0.098093]$	0.057929	$[0.024459, 0.090059]$
Signal: topic concentration	0.024253	$[0.003776, 0.029774]$	–	–
Decisive: USA reference	−0.006054	$[-0.008018, -0.004629]$	−0.010969	$[-0.012416, -0.009411]$
Decisive: log bilateral trade	0.000065	$[0.000006, 0.000120]$	–	–
Decisive: recipient trade dependence	–	–	0.000905	$[0.000829, 0.000982]$
Decisive: diplomatic representation	0.001133	$[0.001022, 0.001294]$	0.000153	$[0.000118, 0.000187]$
Decisive: defense pact	–	–	0.005080	$[0.004717, 0.005396]$
Abstention: USA reference	−0.001461	$[-0.003612, 0.000583]$	−0.008423	$[-0.011137, -0.006136]$
Abstention: defense pact	–	–	0.004746	$[0.004401, 0.005069]$
Abstention: USSR × defense pact	–	–	0.050000	$[0.046020, 0.050000]$

Notes: All coefficients are in original units; a dash denotes an excluded regressor. The main TBIP intervals use 480 accepted of 500 combined roll-call-text TBIP posterior-draw and roll-call bootstrap replications with original neutral starts. GD intervals use all 500 accepted GD posterior-draw and roll-call bootstrap replications with fitted-MLE starts. Both report 2.5th and 97.5th percentiles. The models differ in preference inputs, bilateral covariates, and signal/prior controls, as well as bootstrap initialization; this comparison does not isolate the corpus change. Log bilateral trade and recipient trade dependence have different definitions and scales. Negative signal-scale coefficients imply lower private-signal standard deviations. Bilateral coefficients are components of complete links, not voting-probability effects. The GD USSR–defense-pact abstention coefficient is bound-constrained.

(68.8%). Its percentile interval ends at that bound and describes a constrained estimator; it does not resolve sensitivity to the constraint.

F Exploratory Counterfactual Voting Outcomes

This appendix presents exploratory counterfactual exercises using the 109-parameter ROC specification with P5-specific sender deviations in both influence channels, reported in Section . The results are conditional on this fitted specification and remain provisional pending sensitivity and estimation-uncertainty analysis. The benchmark simulates new votes in the observed order. The no-alliance scenario removes both bilateral channels; the no-learning scenario sets $\kappa_L = 0$; and simultaneous voting removes both mechanisms. Private information, preferences, prior beliefs, and the baseline abstention wedge remain fixed. Each scenario regenerates voting histories using common private-signal shocks, with 100 draws per state and roll call. Outcomes under states zero and one are weighted by $1-\widehat{\rho}_t$ and $\widehat{\rho}_t$, respectively, and then averaged equally across all 2,679 roll calls.

Table 19 reports vote shares and probabilities of meeting simple-majority ($Y > N$) and two-thirds ($Y+N > 0$ and $3Y \geq 2(Y+N)$) thresholds, where Y and N are Yes and No counts. These descriptive thresholds exclude abstentions and need not match each resolution's legal adoption rule. Thresholds are evaluated within each simulation before averaging.

Table 19: Exploratory counterfactual outcomes under the P5-sender specification.

	Yes	No	Abstain	Simple majority	Two-thirds
Panel A: Observed and simulated baseline levels (percent)					
Observed outcomes	55.49	24.32	20.19	73.16	58.34
Simulated observed-order baseline	57.52	23.56	18.92	75.96	63.61
Panel B: Changes from simulated baseline (percentage points)					
No alliance	-2.91	+5.01	-2.10	-1.36	-10.33
	(0.019)	(0.016)	(0.011)	(0.047)	(0.053)
No learning	+2.31	-3.86	+1.55	+12.78	+9.25
	(0.021)	(0.018)	(0.012)	(0.050)	(0.054)
Simultaneous	-1.72	+2.52	-0.80	+16.84	-5.55
	(0.029)	(0.023)	(0.013)	(0.054)	(0.058)

Notes: Results use the ROC model with P5 sender deviations in both channels and 100 private-signal draws per state for each of 2,679 roll calls. State uncertainty is integrated using the fitted prior. Parentheses report Monte Carlo standard errors of paired changes, in percentage points; these measure simulation error and exclude parameter and first-stage estimation uncertainty. Panel A's observed threshold entries are fractions of actual roll calls meeting each threshold; simulated entries are mean probabilities.

Removing both bilateral channels lowers Yes share by 2.91 percentage points, raises No share by 5.01 points, and lowers Abstain share by 2.10 points. The probability of meeting the two-thirds threshold falls by 10.33 points. Removing learning instead raises Yes share by 2.31 points, simple-majority probability by 12.78 points, and two-thirds probability by 9.25 points. Positive informational updating therefore need not increase affirmative support; its direction depends on earlier votes. Under simultaneous voting, Yes share falls by 1.72 points, while simple-majority probability rises by 16.84 points and two-thirds probability falls by 5.55 points. These are different thresholds of the simulated vote-count distribution, so their probabilities need not move in the same direction. The interventions are not additive because each changes the subsequent voting history.

These magnitudes are conditional on the fitted P5-sender specification. The simulated baseline overpredicts the observed simple-majority and two-thirds frequencies by approximately 2.80 and 5.26 percentage points, respectively. Monte Carlo standard errors measure simulation precision; they exclude parameter and first-stage estimation uncertainty. The exercises hold structural parameters fixed under alternative institutions and do not establish welfare improvements or robustness across specifications.

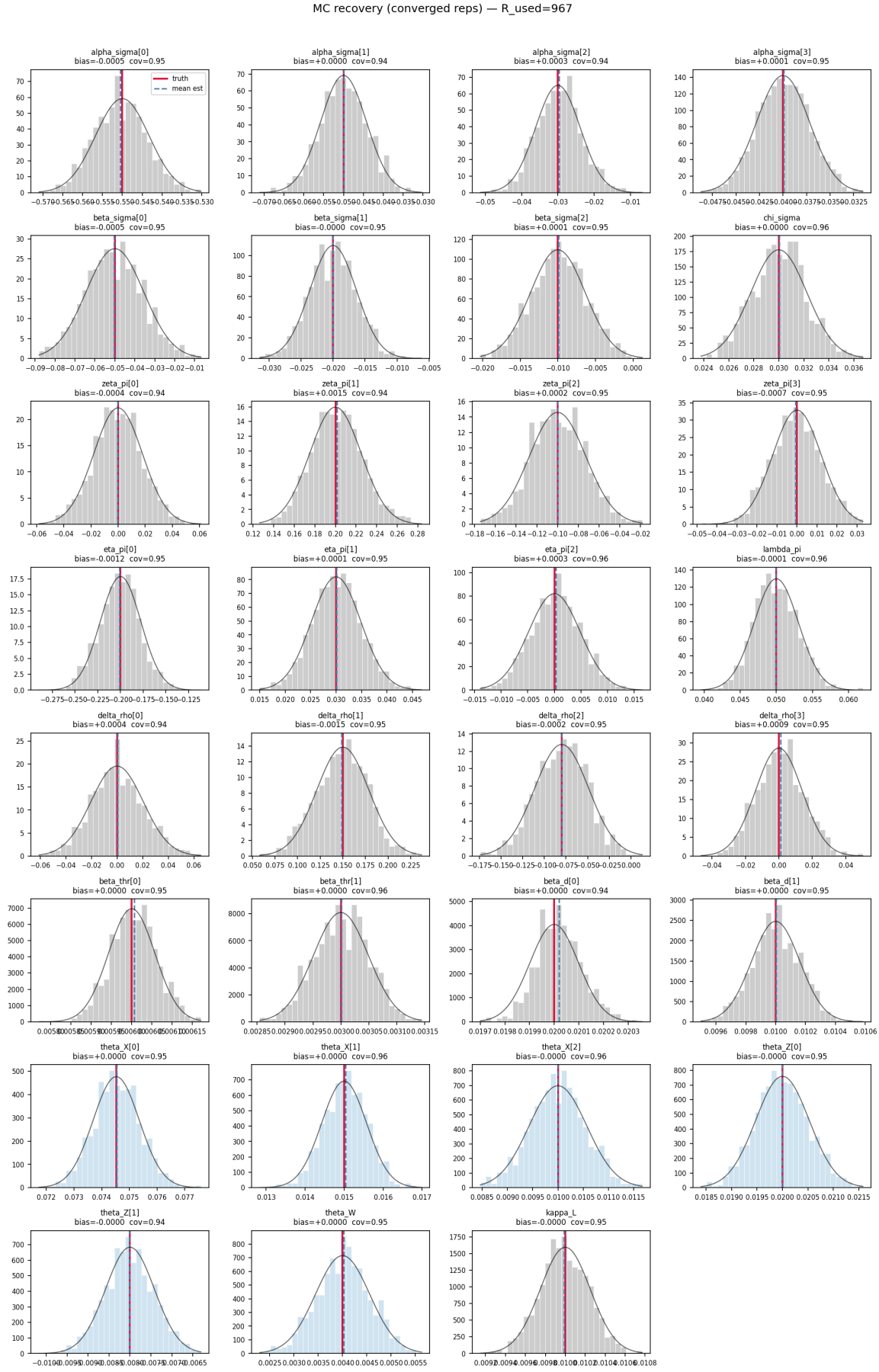


Figure 6: Maximum-likelihood recovery. Sampling distribution of each estimate over $R = 967$ converged replications ($N = 100$, $T = 3000$); the solid line marks the true value and the dashed line the mean estimate. Panel titles report the mean bias and the empirical 95% Wald coverage.

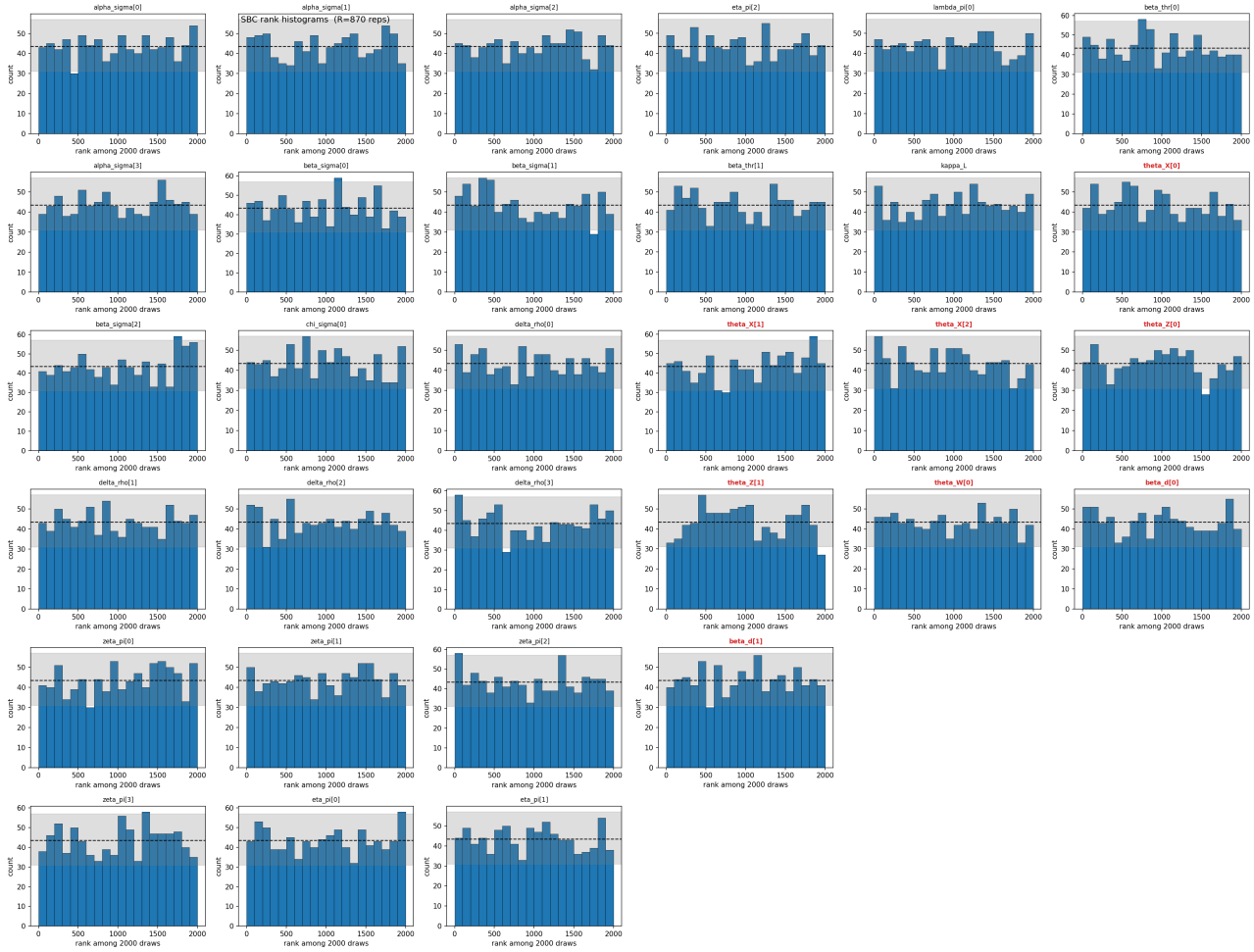


Figure 7: Simulation-based calibration rank histograms for each of the 31 parameters ($R = 870$ converged replications, ranks among $L = 2000$ posterior draws). The shaded band is the interval expected under uniformity; ranks are consistent with uniformity throughout.