

Social Capital, Institutional Rules, and Constitutional Amendment Rates

William D. Blake* Joseph Francesco Cozza[†] David A. Armstrong II[‡]
Amanda Friesen[§]

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Abstract

Why are some constitutions amended more frequently than others? The literature provides few clear answers, as some scholars focus on institutional factors, while others emphasize amendment culture. We bridge this divide with new theoretical and empirical insights. Using data from democratic constitutions worldwide and US state constitutions, we examine how social capital reduces the transaction costs imposed by amendment rules. The results indicate that constitutional rigidity decreases amendment frequency, but group membership, civic activism, and political trust can offset the effect of amendment rules. Our findings have important implications for scholars in public law, constitutional and democratic theory, and social movements.

Keywords: constitutional amendment rates, social capital, constitutional rigidity, transaction costs, US state constitutions

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* Associate Professor of Political Science, University of Maryland, Baltimore County, ORCID 0000-0002-0538-6104, wblake@umbc.edu. The authors wish to thank Lawrence Baum, Adam Brown, Carolyn Forestiere, Hans Hacker, Ran Hirschl, Alexander Hudson, Aubrey Jackson Soller, Susan Sterett, Fernando Tormos-Aponte, Lydia Tiede, Mila Versteeg, Christopher Wleizen, the students of the UMBC Honors College, and the anonymous reviewers for their invaluable feedback. Previous versions of this paper were presented at the Global Law & Politics Workshop and the annual meetings of the Midwest Political Science Association and the American Political Science Association. Replication data can be found at <https://doi.org/10.7910/DVN/ZHIU9M>.

[†] Visiting Assistant Professor, Rice University, ORCID 0000-0001-6508-4282, joseph.cozza@rice.edu.

[‡] Canada Research Chair in Political Methodology and Associate Professor of Political Science, Western University, ORCID 0000-0002-9358-2489, dave.armstrong@uwo.ca

[§] Associate Professor of Political Science, Western University, ORCID 0000-0002-9743-5742, afries4@uwo.ca.

Since the days of Madison and Jefferson, political scientists and legal theorists have questioned whether shorter, entrenched constitutions are more desirable than longer, frequently-amended documents. Behind this normative debate lies an empirical question: what factors shape amendment rates? If constitutional designers desire entrenchment, for example, what should they do? According to some scholars, rigid amendment rules constrain elites seeking constitutional change (Dixon and Holden 2012; Elkins, Ginsburg, and Melton 2009; Ferejohn 1997; La Porta et al. 2004; Lijphart 2012; Lorenz 2005; Lutz 1994; Negretto 2012; Rasch and Congleton 2006; Tsebelis 2022) while others have suggested amendment culture plays a larger role than institutional rules (Contiades and Fotiadou 2013; Fruhstorfer and Hein 2016; Ginsburg and Melton 2015; Tarabar and Young 2021).

Existing accounts of constitutional amendments have major weaknesses. Institutional approaches analyze how rules affect the supply of amendments without examining the conditions that stimulate amendment demand. Meanwhile, a nation's amendment culture may be endogenous to its amendment rules by creating norms about the desirability of amendments. Thus, it is unclear whether amendment culture changes over time and, if so, why. And no study to date has disaggregated the behavior of different players in the amendment process.

We advance the study of constitutional amendments by modeling amendment rates as an interplay between social forces and institutional rules. We theorize that social capital creates a political environment more conducive to constitutional reform. Robust levels of social capital help elites, citizens, and social movements overcome the transaction costs created by amendment rules. We follow Keele's (2007, 242) definition: "Social capital is comprised of civic participation and trusting attitudes." Thus, our analysis leverages the bundle of attitudes and behaviors scholars use to operationalize the effects of interpersonal networks: political trust, civic engagement, and group membership (Putnam 2000, 19). Our theory is not contingent on resolving scholarly disagreements about the formation and perpetuation of social capital. While these debates are important, it is the existence of trusting attitudes

and civic activity that animates our analysis.

We analyze amendment rates of democratic constitutions globally and US state constitutions, using social capital indicators from three datasets. Our cross-sectional and longitudinal analyses indicate that amendment rules matter, but different polities are better equipped to navigate amendment rules depending on their civic connectedness. The American states provide a rigorous test for our theory: sub-national amendments occur frequently, suggesting transaction costs are low. Nevertheless, the results of a novel, two-stage model of amendment proposals and ratifications indicates that social capital trends affect state-level constitutional reform.

1 Determinants of Amendment Rates

Scholars have not discovered a clear relationship between amendment rules and amendment rates partly because operationalizing amendment processes is difficult. Many studies make different assumptions when constructing ordinal or linear indices, creating weak cross-measure correlations (Ginsburg and Melton 2015, 698). A more sophisticated measure comes from Tsebelis (2022), who finds a heteroskedastic relationship, as flexible amendment rules provide a necessary but insufficient condition for constitutional change. These methodological challenges have produced contradictory findings. Some studies suggest that rigid amendment rules reduce amendment frequency (Dixon and Holden 2012; Fruhstorfer and Hein 2021; Lutz 1994; Negretto 2012; Lijphart 2012; Tsebelis 2022), while others produced null results (Ferejohn 1997; Ginsburg and Melton 2015; Lorenz 2005; Rasch and Congleton 2006).

As legal theorist Richard Albert (2019, 2, 97-98) acknowledges, “No part of a constitution is more important than the rules we use to change it,” yet “studies of formal amendment difficulty are insufficient and incorrect if they evaluate only the codified rules of change without attention to the larger political and cultural contexts in which those rules are situated.” One prominent amendment culture study comes from Ginsburg and Melton (2015), who,

in a break from their previous scholarship (Elkins, Ginsburg, and Melton 2009), find that institutional factors do not influence amendment rates when models include a measure of amendment culture. However, their measure does not reflect attitudinal data, even though Ginsburg and Melton (2015, 699) define amendment culture as a “shared set of attitudes about the desirability of amendment.” Instead, they measure culture using the amendment rate of a country’s prior constitution. This is problematic considering amendment rules often do not change when a nation replaces its constitution (Miller et al. 2021, 16).

Tarabar and Young (2021) use Hofstede cultural dimensions as proxies for constitutional attitudes and find that more individualistic and long-term-oriented societies have higher amendment rates, while cultures that avoid uncertainty have fewer amendments. The Hofstede measures are based on survey data collected from employees of IBM subsidiaries during the 1960s-70s.¹ Even if these data were more extensive and nationally representative, they are based on an assumption that culture does not vary over time or sub-nationally.

Social capital offers an opportunity for theoretical and empirical insights beyond what the literature offers. While social capital does not directly measure amendment attitudes, trust and activism may be factors that shape a constitutional culture. Scholars have found that social capital affects policy innovation (Putnam 1993, 82-120; 2000, 346-47), and we extend this insight to constitutional innovation. Constitutional rules impose transaction costs (Buchanan and Tullock 1965), and, in general, social capital reduces transaction costs (Fukuyama 1995). Thus, we predict amendments will be adopted more frequently in polities with higher levels of social capital or at times within a polity’s history when social capital is comparatively higher.

1.1 Amendment Rules and Transaction Costs

Amendment rules structure negotiation and decision costs to make constitutional reform less frequent than legislative change. Constitutional designers commonly employ several

¹ See the online Appendices, pp. 20-22, for an extended critique.

strategies. Requiring supermajorities to approve amendments means any change must satisfy the preferences of a larger proportion of the decision-making body. Increasing the number of veto points in the amendment process serves a similar function. Some constitutions provide multiple amendment paths, increasing the strategic choices available to reformers. Social movements and interest groups are sensitive to these costs, taking positions on the rigidity of amendment rules based on whether they prefer the status quo (Dinan 2016).

Higher decision costs creates higher renegotiation costs, especially when constitutions fail to specify important details (Dixon and Ginsburg 2011). Rigid amendment rules raise information costs because decision-makers must be confident that any lasting bargain they reach will be beneficial. When information costs associated with constitutional change are high, the familiarity of the status quo may seem preferable (Blake and Anson 2020; Zink and Dawes 2016). Flexible amendment rules, by contrast, facilitate constitutional experimentation because if a reform causes problems, they can be more easily solved with a future amendment.

Elites are better at navigating these costs than citizens because they possess strategic incentives and clearer preferences (Fruhstorfer and Hein 2021; Negretto and Sánchez-Talanquer 2021). When negotiation and decision costs are high, political actors may utilize other strategies. Judicial review can alter existing constitutional provisions, creating a need for public trust in the courts. Elites can also pursue state-level constitutional reform or entrench constitutional principles in ordinary legislation (Woodward-Burns 2021).

1.2 Social Capital and Transaction Costs

The trust, reciprocity, and civic activity produced by social capital help elites, ordinary citizens, and social movements overcome the transaction costs identified earlier. Social capital reduces coordination costs for social movements, facilitating member recruitment, resource mobilization, and coalition formation (Edwards 2013). With greater resources, social movements can pursue costlier reform strategies. For example, as social capital in the United

States increased in the early 20th century (Putnam 2000), temperance groups updated their strategies from enforcing local nuisance laws to state-level constitutional reform, followed by a federal statute banning the interstate shipment of alcohol, then the Eighteenth Amendment (Compton 2014). Social movements reduce information costs for elites by clarifying constituent preferences (Tavits 2006). When groups are active on both sides of an issue, social movements illuminate the compliance costs of a potentially polarizing amendment.

Social capital facilitates elite coalitions needed to satisfy amendment rules. Boix and Posner (1998) draw connections between social capital and consociational politics, where cross-cleavage/cross-party coalitions are common. Trust mitigates the risk aversion that comes with changing policies, and trust increases the openness of political agendas, helping elites overcome status quo bias. Diani (2006, 134) argued that social capital is “conducive to more open attitudes by the elites towards political challengers.” Social capital generates democratic deliberation, which diminishes political polarization (Fishkin 2011). Distrust, however, is associated with the rise of populism and partisan exploitation of constitutions (Landau 2013).

Evaluating how citizens approach constitutional reform is also important, as referenda are the most common amendment ratification method (Elkins and Hudson 2019). On ballot measures, voters utilize elite cues as heuristics (Bowler and Donovan 2000). For those cues to work, they must seem trust-worthy. Indeed, political trust strongly affects support for referenda in Switzerland (Kriesi 2005). Social capital enables social movements to provide more information to voters. If citizens have no direct role in the amendment process, they can still hold elites accountable. Claibourn and Martin (2007) find that participation in voluntary societies is associated with increases in political knowledge and accountability. Furthermore, Uslaner (2000, 570) argued that trusting voters “place a high value on compromise and legislative productivity, rather than ideological purity and stalemate.”

Nevertheless, the influence of social capital may be limited. Putnam and others assume that when citizens join groups, they form social bonds that translate into trusting attitudes.

However, voluntary memberships do not automatically stimulate social trust, and expressions of trust vary across citizens—usually “winners” in society are more trusting (Newton 2001). Social capital trends can also obscure or exacerbate racial inequalities (Hero 2003), and the effect of social capital on political involvement is gendered (Lowndes 2004).

Building on Tsebelis’ (2022) finding that constitutional rigidity has heteroskedastic effects, we predict social capital provides a necessary but insufficient condition for amendments. That is, low social capital levels are strongly associated with fewer amendments, but the effect of robust social capital is more variable. Amendment frequency depends on a number of contextual factors, but these factors might only come into play in trusting and participatory political environments.

2 Analysis

We analyze an aggregate, cross-national dataset of amendment rates using World Values Survey (WVS) data; a longitudinal, cross-national dataset of amendments using V-Dem data; and a longitudinal, US state-level dataset of amendment proposals and ratifications. We limit our analysis to democratic nations, where amendment rules are more likely to constrain regimes, and social capital is more likely to influence political action. Following Tsebelis (2022), we define democracies as nations with a Polity IV score of six or higher. In the WVS and US state analyses, we utilize distributional negative binomial regression. In the V-Dem analysis, the dependent variable indicates the adoption of at least one amendment in a country-year, so we use logistic regression.²

We use Bayesian simulation to estimate all models (Western and Jackman 1994). Gelman and Hill (2007) suggests that Bayesian analysis recovers variance components more reliably and permits the evaluation of uncertainty around random effects estimates. We characterize the uncertainty of model parameters and predictions by summarising the posterior distribu-

² These models differ from the linear models utilized in previous research. See the Appendices, pp. 9-11, 20-22.

tion. What scholars would call “confidence” in the Frequentist paradigm is analogous to a Bayesian’s “credibility.”

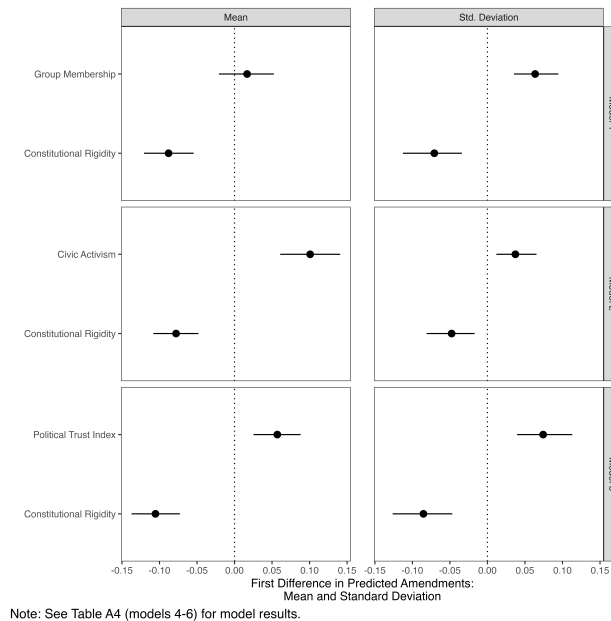
To measure credibility, we calculate the proportion of the posterior distribution that lies on the same side of zero as the posterior mean. We refer to this as the Bayesian p-value or the “posterior probability.” Credible intervals of 80% allow for hypothesis testing at the 90% Bayesian p-value level. We use an average first-difference approach to assess how constitutional rigidity and social capital affect the expected value and residual variability in amendment rates. Holding other variables at their observed values, we measure $\pm\delta = \frac{1}{2}\text{SD}(X)$ of the variable of interest and average the difference in predicted values for $x - \delta$ and $x + \delta$.

The WVS provides three indicators of political trust: confidence in government, political parties, and courts. A reliability analysis suggests that a single index represents the main source of variation among these variables (Cronbach’s $\alpha = 0.92$). We also create additive indices measuring participation in voluntary organizations and civic activism—based on respondents’ petitioning, boycotting, and protest history. While these variables do not directly measure resources available to social movements, scholars use these data to compare the effects of social movements across nations (e.g., Norris, Walgrave, and Van Aelst 2005). Each indicator is a weighted, cross-wave national average of the 57 democracies surveyed by the WVS.³

Here, the dependent variable is the number of years in which amendments were ratified during the democratic lifespan of the current constitution, with the logged democratic age of the constitution serving as an exposure term. We leverage Tsebelis’ (2022) constitutional rigidity measure—a sum of each approval threshold in the amendment process, accounting for inter-chamber differences in party composition. Previous studies indicate amendment rates are higher in countries with longer constitutions (e.g., Ginsburg and Melton 2015). Thus, our models control for constitutional length, as measured by the logged number of

³ See the Appendices, pp. 1-3, for more details.

Figure 1: Effects of Constitutional Rigidity and Social Capital on Amendment Rates



words.

Figure 1 shows the effect of rigidity and social capital measures from three models on the amendment rate (mean function) and the residual standard deviation. A standard-deviation change in constitutional rigidity tends to reliably decrease the number of amendments ratified over time. All differences are credible at the 90% level. A similar change in social capital reliably increases the number of amendments ratified for civic activism and the political trust index, while the effect of group membership is not credibly different from zero.

Social capital also tends to increase the residual variation of amendments, supporting our hypothesis that social capital is a necessary but insufficient condition. In other words, as social capital increases, so does the possibility of adopting amendments, but other country-specific factors determine whether amendments become a reality. Consistent with Tsebelis (2022), more rigid constitutions have credibly less residual variation.

The Indian constitution has been amended over 100 times since 1950. India has relatively flexible amendment rules—in the 10th percentile according to Tsebelis (2022)—as many provisions can be amended by a parliamentary majority. But, as we have argued, flexibility

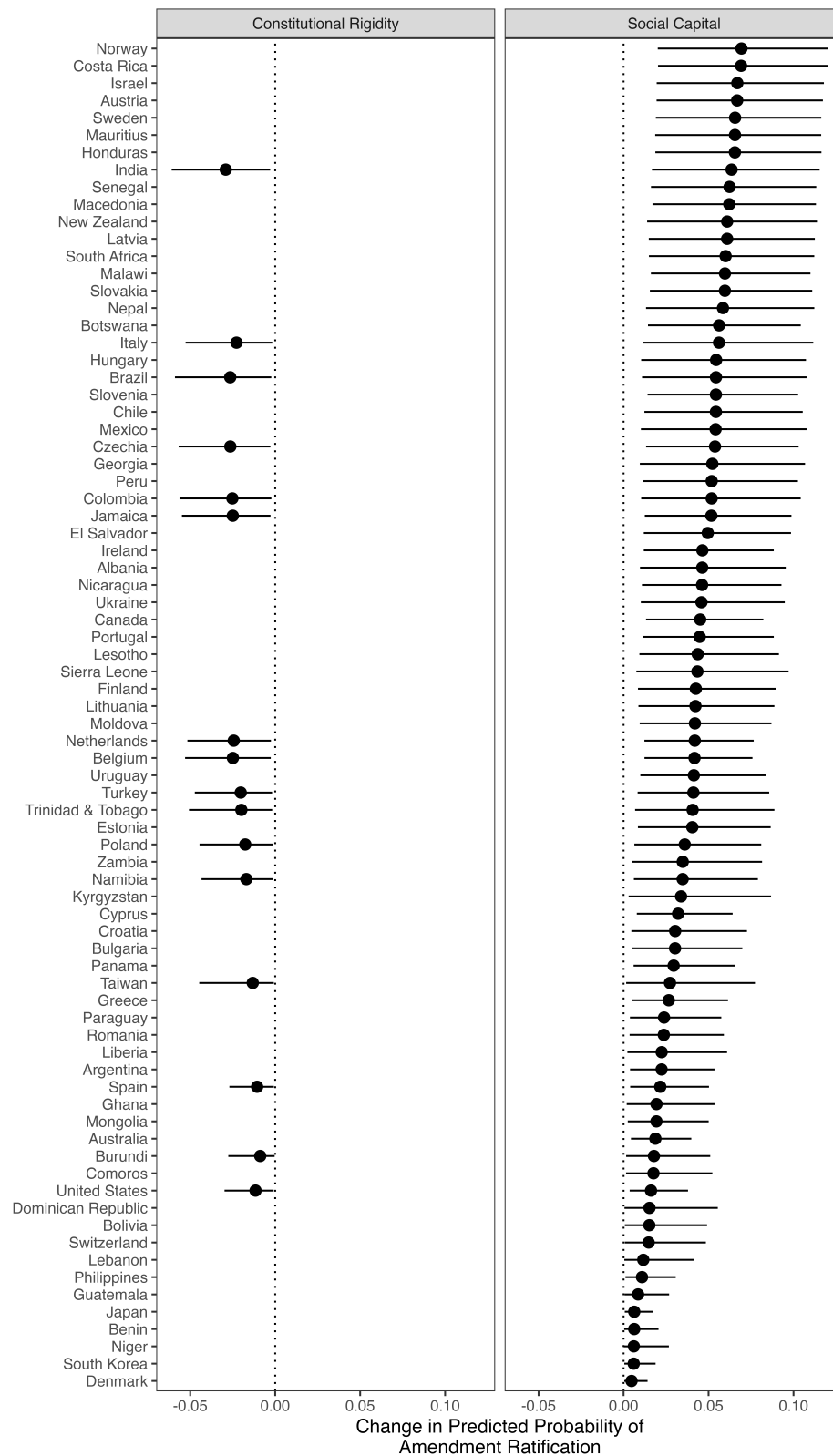
does not guarantee amendment frequency. Social capital provides another important factor: India ranks among the top half of countries surveyed in the 2014 WVS with respect to confidence in government (22/59) and the top third in confidence in courts (20/59) and political parties (15/57). Indian citizens are also highly engaged, socially and politically. India ranks among the top 10 countries in group membership (out of 57) and 11th (out of 54) for attending political demonstrations. By contrast, Spain has a rigidity score similar to India (13th percentile), yet India's amendment rate is more than ten times higher. One reason might be that Spain ranks 55th in confidence in government, 44th for courts, 46th for parties and 35th for group membership, only outranking India in participation in demonstrations (second).

To examine within-country, over-time variation, we analyze V-Dem data on 80 democracies from the adoption of their current constitution through 2013. We replicate Tsebelis' (2022) rigidity measure using legislative fractionalization, courtesy of the PolCon dataset, as a substitute for the annual inter-chamber distance in party composition. The measures correlate at around 0.7 in 2013, although there is only within-country variation in sixteen nations. We pair this data with V-Dem's Civil Society Participation Index (CSPI), an annual, composite indicator of the degree to which organizations are engaged in politics.

Like Beck, Katz, and Tucker (1998), we estimate a binary time-series cross-sectional model, although we treat the data as hierarchically structured—with country random intercepts—and model time dependence by estimating a random effect on the spell length. Like Tarabar and Young (2021), we control for real per capita GDP (growth rate and logged base level), Polity's executive constraint scores, episodes of political violence, and ethnic fractionalization. Following Bell and Jones (2015), we create between- and within-country variables. We are primarily interested in within-country variation, as the WVS analysis explored between-country variation.⁴

⁴ See the Appendices, pp. 3, 11-17, for more details.

Figure 2: Effects of Constitutional Rigidity and Social Capital on Amendment Probability in Democracies



See Table A6 (model 1) for model results.

The left portion of Figure 2 visualizes the credible, negative effect of constitutional rigidity in all sixteen nations where constitutional rigidity varies over time. This finding suggests that as divisions between legislative chambers increase, the probability of amendment decreases. As indicated on the right side of Figure 2, higher CSPI scores are associated with reliably higher amendment activity in every democracy save two. The magnitude of these effects vary across countries, which is unsurprising given the heteroskedastic findings from the WVS analysis.

Moreover, the variable effect sizes are consistent with existing scholarship. According to (Ginsburg and Melton 2015, 689), India, Brazil, and Mexico are paradigmatic examples of flexible amendment cultures, while the US and Japan represent rigid amendment cultures. As displayed on the right side of Figure 2, social capital exerts a comparatively large effect in India (eighth-largest), Mexico (22nd largest) and Brazil (24th-largest), and a smaller effect in the US (67th-largest) and Japan (75th-largest). Thus, our results suggest that civil society plays an important role in shaping amendment culture.

It is possible that social capital trends are endogenous to constitutional reform. Some democratically-transformative amendments may enable activism or inspire trust. The structure of the V-Dem dataset allows us to test for endogeneity. As detailed in the Appendices (pp. 4-7), there is no evidence that constitutional amendments affect future social capital levels. We also examine whether social capital exerts stronger effects when citizens are directly involved in the amendment process. Fifty-nine of the 78 democracies in the analysis hold amendment referenda, but within-country social capital fluctuations produce effects that are statistically no different from non-referenda nations.⁵

Amendment rules often require sequential decision-making, yet scholars have created singular measures of constitutional rigidity, rather than one for each player in the sequence. We fill this gap in the literature with data on US state constitutional amendments. In every state but Delaware, voters ratify amendments proposed by legislatures, and we model

⁵ See Appendix Table A7.

these actions separately.⁶ We replicate Tsebelis' (2022) constitutional rigidity measure but make separate calculations for the proposal and ratification stages. We use a moving average of Hawes, Rocha, and Meier's (2013) annual, state-level social capital indicator—a factor analysis combining measures of group membership, civic engagement, and community volunteerism from every continental state. Our data are organized biennially from 1987 to 2009.

The amendment proposal model includes within- and between-state levels of social capital and rigidity, while controlling for word length. Like the V-Dem analysis, we include state-level random intercepts to capture unobserved unit-level factors that affect amendment demand (like Alabama's home rule constraints).⁷ We also model ratification as a function of within-state social capital changes, adding controls for the between-state effects of rigidity and social capital.⁸ We then simulate the changes in amendment proposals, given a half-standard-deviation plus-or-minus change in social capital. Next, we predict the number of ratified amendments, based on a similar change in social capital *and* accounting for the change in proposed amendments.⁹

Within the structure of American federalism, state constitutions are more detailed and dynamic than the US Constitution. The transaction costs imposed by state constitutional rules seem manageable, given the frequency of change. Indeed, we find over-time changes to a legislature's party composition do not credibly affect amendment proposal rates. Likewise, between-state rigidity differences (e.g., a two-thirds legislative vote versus a simple majority) do not affect amendment proposals, although between-state differences in referendum rules credibly affect voter ratification rates. Meanwhile, within-state increases in social capital

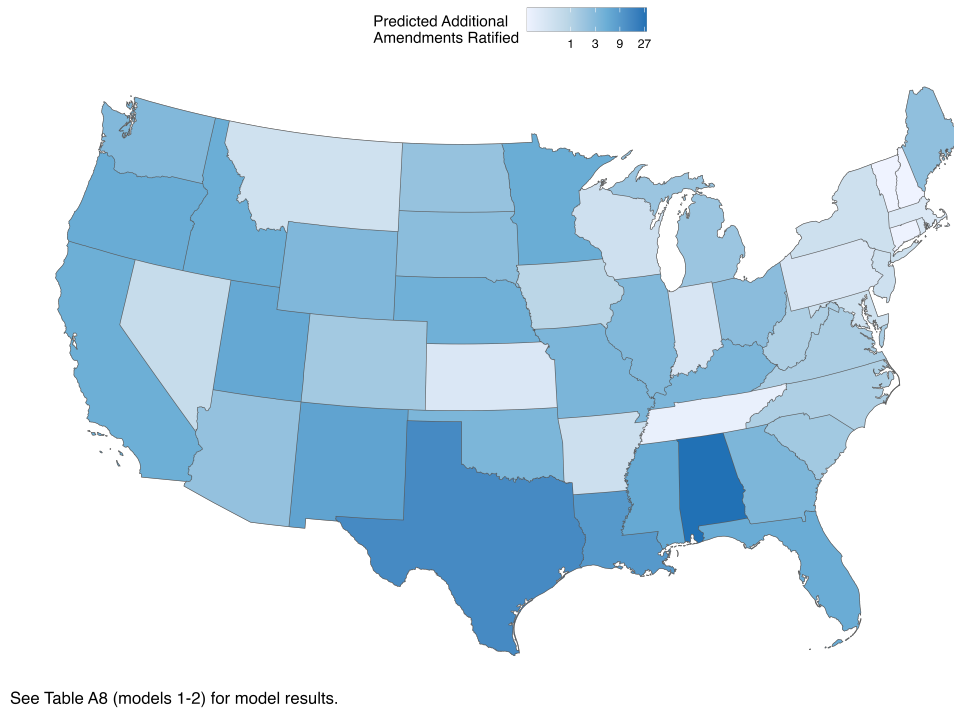
⁶ We exclude amendments proposed via the constitutional initiative.

⁷ See <http://encyclopediaofalabama.org/article/h-1153>.

⁸ Only Florida changed its referendum rules during this timeframe, so we omit within-state rigidity differences from the ratification model.

⁹ See the Appendices, pp. 3-4, 17-20 for more details.

Figure 3: Effect of Social Capital on State Constitutional Amendment Ratifications



reliably produce a higher number of proposed amendments, and this increase translates more or less directly into additional ratifications through an increased offset term in the ratification model.

As visualized in Figure 3, social capital helps to explain cross-state differences in amendment activity, producing the largest effect in Alabama—the state which adopted the most amendments. In Massachusetts, which had the fewest amendments, social capital exerted the 40th-largest effect. Using the posterior mean of predicted amendment proposals as the offset term, social capital effects are statistically reliable in 32 of the 47 states analyzed.

3 Discussion

Our analyses of constitutions in US states and democracies globally indicate that amendment frequency is a product of amendment rules, group membership, civic activism, and political trust. Unlike previous studies, we demonstrate how these factors affect constitutional

reform over time and across the stages of the amendment process. We also build upon prior research that finds social capital facilitates social movement organizing and elite coalition formation. Consequently, we contribute to ongoing debates regarding the umbrella term of social capital by specifying how specific attitudinal and behavioral components work and why they matter.

At the same time, this study has limitations and leaves other important questions undressed. Data constraints prevent testing which social capital indicators play a bigger role at which stage of the amendment process. While social norms, like trust and participation, are factors that influence amendment culture, collecting direct, comparative measures of constitutional attitudes would be helpful. It remains to be seen if our theory generalizes to other contexts, including the erosion of democratic constitutions, constitutions in non-democracies, or international legal agreements (like Brexit). It is also unclear whether our theory applies when countries replace, rather than amend, their constitution.

Experiments and case studies would provide further insights into how the causal mechanisms operate and vary across contexts. Such research would provide greater precision to the general pattern we identify—that higher levels of social capital and less rigid amendment rules may increase amendment activity, but these effects are variable. On the other hand, when social capital is low and rigidity is high, constitutional change often becomes extremely difficult.

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5 Online Appendices: Social Capital, Institutional Rules, and Constitutional Amendment Rates

In the Appendices, we describe the data, detail our estimation strategy, display numerical results, and undertake some robustness analyses.

5.1 Data

5.1.1 Cross-national Data

Like Tsebelis (2022), we include Israel in the analysis, but we do not include the United Kingdom because its uncoded constitution renders it an exceptional case. There is debate in the literature as to when the current UK constitution was enacted, which affects the measurement of its amendment rate and length (cf. Tsebelis 2017, 2022)—a dependent variable and control in the analysis.

We also use Tsebelis’ measure of constitutional rigidity, which is a sum of the approval thresholds of the institutions required to consent to a constitutional amendment based on data provided by the Constitute Project. When amendments require the approval of a bicameral legislature, the measure is based on the Euclidean distance between the partisan composition of the two chambers. That is, if one legislative chamber is composed of parties with proportions $x_1, x_2, x_3, \dots, x_n$, and a second legislative chamber is composed of parties with proportions $x'_1, x'_2, x'_3, \dots, x'_n$, then the distance between the two chambers is calculated as $\sqrt{(x_1 - x'_1)^2 + (x_2 - x'_2)^2 + (x_3 - x'_3)^2 + \dots + (x_n - x'_n)^2}$. The index also includes adding or subtracting a small constant for other procedural requirements that affect rigidity, e.g., a quorum requirement or a mandate that an amendment be passed twice in the same legislative session.

The group membership variable is the total number of group types in which a respondent has either an active or inactive membership. We limit the total number of group types to subjects of inquiry on every iteration of the WVS: church or religious organizations; sport or recreational organizations; art, music or educational organizations; labour unions; political parties, environmental organizations; professional organizations; humanitarian or charitable organizations; and other organizations.

Table A1: Social Capital Indicator Correlation Matrix

	Government Confidence	Party Confidence	Court Confidence	Group Membership
Party Confidence	0.873			
Court Confidence	0.700	0.767		
Group Membership	0.368	0.318	0.261	
Civic Activism	-0.048	0.047	0.269	0.263

Consistent with other scholarship (Welzel 2013), we limit the civic activism variable to petitioning, boycotting, and protesting, as these activities are the only ones that appear on every iteration of the WVS. As displayed below, one response option given to survey participants is that they “might do” a civic activity. For our index, however, we only count instances when a respondent has engaged in a given activity. As displayed in Table A1, some of these indicators are strongly correlated (such as Government Confidence and Party Confidence), while others are not (like Government Confidence and Civic Activism).

Table A2: Correlation All WVS Data and Pre-2014 Data

Variable	Correlation
Government Confidence	0.959
Party Confidence	0.982
Court Confidence	0.985
Group Membership	0.938
Civic Activism	0.995

Data for the dependent variable end in 2013, but we use all available WVS data, including surveys since 2013. As discussed in the body of the paper, each social capital indicator is constructed by creating a country-wave average and then a cross-wave, national average for countries surveyed in multiple waves. We include all available WVS data for several reasons. First, it provides additional degrees of freedom for a small sample and broadens the diversity of democratic nations included in the analysis. Second, the dependent variable measures the amendment rate for each nation’s current constitution, and no democratic nations have replaced their constitutions since then. Finally, as indicated in Table A2, there is a very strong correlation between variables constructed using all available data and pre-2014 data.

The wording for all the WVS confidence questions is as follows:

- I am going to name a number of organizations. For each one, could you tell me how much confidence you have in them: is it a great deal of confidence, quite a lot of confidence, not very much confidence or none at all?
 - The government (in your nation’s capital)
 - Political parties
 - The courts

- Now I am going to read out a list of voluntary organizations; for each one, could you tell me whether you are a member, an active member, an inactive member or not a member of that type of organization?
 - Church or religious organization
 - Sport or recreational organization, football/baseball/rugby team
 - Art, music or educational organization
 - Labour union
 - Political party
 - Environmental organization
 - Professional organization
 - Humanitarian or charitable organization
 - Other organization
- Now I'd like you to look at this card. I'm going to read out some different forms of political action that people can take, and I'd like you to tell me, for each one, whether you have actually done any of these things, whether you might do it or would never, under any circumstances, do it.
 - Signing a petition
 - Joining in boycotts
 - Attending peaceful demonstrations

5.1.2 Cross-national Longitudinal Data

The Civil Society Participation Index, courtesy of Varieties of Democracy (V-Dem), represents an aggregation of four indicators measuring the degree to which organizations engage in politics. The first indicator measures the participatory environment in which civil society organizations (CSOs) operate, with potential scores ranging from zero, where the state sponsors all organizations, to three for societies with diverse CSOs in which citizens at least occasionally participate. Next, the index includes an indicator that gauges the extent to which rulers routinely consult CSOs. The degree to which women can participate in CSOs forms another dimension of the index. Finally, the index includes a measure of the centralization of candidate nominations (Bernhard et al. 2017, 347-48).

The Political Constraint (PolCon) dataset was developed by Henisz (2000). His measure of legislative fractionalization is equal to the probability that two random draws from one legislative chamber are from different parties. Given the larger N , we include additional controls. Courtesy of Penn World Table, we control for real logged per-capita GDP and real GDP annual growth. Like Tarrabar and Young (2021), we also control for the Executive Constraint component of each nation's Polity IV score and ethnic fractionalization, using the Historical Index of Ethnic Fractionalization (Drazanova 2020), which corresponds to the probability that two randomly drawn individuals within a country are not from the same ethnic group. Finally, we include a dichotomous indicator of a major episode of political violence per country-year, as defined by the Center for Systematic Peace.

5.1.3 State Data

The Book of the States, published by the Council of State Governments, provides data on the number of amendments proposed and ratified in each state, as well as the length of each state's constitution. We use *The Book of the States* and the website Ballotpedia to exclude the number of amendments that were proposed through the constitutional initiative process that exists in 18 states. This allows for a direct comparison across the states of the number of amendments proposed by state legislatures and ratified by voters in Figure A3.

We also use Ballotpedia for information regarding the partisan make up of state legislatures. Following the formula of Tsebelis (2022), we take the Euclidean distance between the two chambers every year before taking a biennial average. Nebraska, which uses non-partisan elections to its single legislative chamber, takes a Euclidean distance value of zero. Like Tsebelis (2022), we add 0.01 if states limit the number of amendments that can be placed on the ballot at one time, if states require an amendment to pass in multiple legislative sessions, if amendments are subject to a quorum requirement, or if a referendum requires a majority of those voting in the election, rather than those voting on the ballot measure (effectively turning an abstention into a no vote). We subtract 0.01 for states that have the constitutional initiative, periodic constitutional convention referendums, and emergency or alternative amendment provisions.

We include information on the referendum process when analyzing amendment proposals because we assume that state legislators will act strategically based on whether they anticipate an amendment will be ratified by the voters. The size of a majority required in a referendum is likely to play a part of this calculation.

Finally, Hawes, Rocha, and Meier (2013) construct their state-level social capital measure using data from the marketing research firm MediaMark Research, Inc (MRI). Every other year, MRI interviews over 20,000 respondents for its publication, *The Survey of the American Consumer*. The resulting index is constructed using a factor analysis of 22 items from the MRI survey, which measure community organizational life, engagement in public affairs, and community volunteerism.

5.2 Addressing Endogeneity

The cross-sectional, longitudinal structure of the V-Dem data provides an opportunity to test whether social capital trends are endogenous to constitutional reform. For high salience amendments that expand democratic rights, the causal arrows may flow in both directions. That is, social capital could mitigate the transaction costs associated with expanding democratic rights, while the creation of new rights then spurs higher levels of political trust and future citizen activism.

Thanks to Tsebelis (2022, 291-92), researchers now have a sense as to which amendments around the world are democratically significant and which are more routine. Tsebelis consulted country experts to classify every amendment event in a country-year into one of three categories: insignificant, significant, or exceptionally significant. Significant amendments “alter (but do not transform) key institutional features of the legislative, executive, or judicial bodies of government (or their relationship); expand the electorate (but not fundamentally alter it) in some way; or enhance individual rights.” The Nineteenth Amendment

to the US Constitution, which expanded the right to vote to women, falls under this category. Meanwhile, exceptionally significant amendments “transform how legislative bargaining or interbranch relations transpire, introduce an entirely new class of individual rights to a citizenry, or were subsequently deemed ‘unconstitutional’ by the country’s Supreme Court.” Ending slavery in the United States via the Thirteenth Amendment qualifies as exceptionally significant.

The subset of “significant” and “exceptionally significant” amendments provide a more fruitful test for an endogenous relationship with social capital than using *all* amendments. Many constitutional amendments are low salience affairs that do not impact the rights of citizens one way or the other. In these instances, there is no theoretical reason to suspect endogeneity. For example, the CSPI score for the United States remained at 0.807 consistently for several years in the 1910s. Following the ratification of the Nineteenth Amendment, the CSPI score immediately jumped up to 0.825 and remained at that level for another decade. On the other hand, the ratification of the 27th Amendment in 1992, which imposes procedural requirements on congressional pay raises, made no impact to the CSPI score of 0.985, which held constant across the 1990s.

To identify and correct for endogeneity, we employ an instrumental variables approach using the Stata package `xtivreg2`. This package assumes that both the dependent variable and potentially endogenous regressor are continuous. In our analysis, the potentially endogenous regressor is dichotomous: whether one or more significant amendments were ratified in a given country-year. Nevertheless, assuming our amendment variable exerts a constant effect, there is no need to use logistic regression in the first stage of the analysis. Clustering the standard errors on each country is sufficient. Conducting logistic regression is theoretically more efficient, but we obtain similar results under either specification.¹

To be effective, instrumental variables should correlate with the potentially endogenous regressor but not the dependent variable. In other words, they need to be associated with constitutional reform but not social capital. For instrumental variables, we use the length of each country’s constitution, measured by the natural log of its word count, and the length of time since the constitution had last been amended. As we explain in the main text, longer constitutions tend to have a greater need for amendments, and we also predict that the length of time a constitution has been stable affects the likelihood of change at a given point in time. Of course, constitutional rigidity would have made for an effective instrumental variable, but there is only significant within-country variation in 16 nations. By excluding it, our sample spans 76 democratic nations.

The goal of this analysis is to evaluate whether the kind of positive, instantaneous effect of amendments like the Nineteenth Amendment on American social capital occurs systematically. Because of the immediate nature of the observed effect, we use a first difference approach in lieu of a fixed effects model, which is preferable when the effect of the independent variable lasts over several years. Both first difference and fixed effects estimators are unbiased and consistent, but they make different assumptions about the error term. The first difference estimator assumes errors are well behaved in differences, while the fixed effects

¹ See <https://www.statalist.org/forums/forum/general-stata-discussion/general/1524185-instrumental-variable-using-panel-data-with-binary-endogenous-variable>.

Table A3: Instrumental Variables Regression Models of Social Capital Trends

Term	(1)	(2)
Amendment Event	0.002 (0.003)	0.004 (0.006)
Freedom of Expression	0.548*** (0.146)	0.548*** (0.146)
Constant	0.002*** (0.001)	0.002*** (0.001)
Weak Instruments (F-Stat)	109.06***	54.913***
Overidentification (Sargan)	0.001	0.016
Endogeneity (WH F-Stat)	0.004	0.014
R^2	0.266	0.266
N	2,500	2,500
Countries	76	76
Amendment Type	Significant/ Exceptionally Significant	Exceptionally Significant

estimator assumes errors are well behaved in levels.

First difference models transform regressors $X_{i,t}$ to $\tilde{X}_{i,t} = X_{i,t} - X_{i,t-1}$ and the dependent variable $y_{i,t}$ to $\tilde{y}_{i,t} = y_{i,t} - y_{i,t-1}$. Thus, it is well-suited to capture the effect of a constitutional amendment, which takes the value of 0 at time $t-1$ and 1 at t , on a change in social capital between time $t-1$ and t .² However, if significant constitutional amendments occur in consecutive years, this model assumes each subsequent year with an amendment would not affect social capital, as the amendment variable takes a value of 1 at time $t-1$ and a value of 1 at time t (leading to a difference of $1-1=0$). Significant constitutional change is rare in general, and it is even rarer to find it occurring in consecutive years (81 times out of an N of 2,500). Our results remain the same whether we include these instances or drop them.

In Table A3, we report the results of two instrumental variable regression models of changes to social capital. The dependent variable is V-Dem's Civil Society Participation Index. The independent variable of interest is the adoption of one or more potentially rights-enhancing amendments. Model 1 covers any amendment deemed significant or exceptionally significant. Model 2 is limited to exceptionally significant amendments only. In both models, we include V-Dem's index measure for freedom of expression and alternative sources of information as a control, as we predict social movements will be more successful in organizing and operating in nations with more open information environments. This index is an annual composite measure of freedom of discussion, academic expression, and cultural expression, government censorship efforts, media bias, and self-censorship.

As predicted, the coefficients of both variables are positive in both models. However, only the Freedom of Expression variable achieves conventional levels of statistical significance.

² Sometimes, the change in social capital does not occur until time t and $t+1$. For example, the CSPI value in the United States increased between 1865 and 1866, the *former* year being the one in which the Thirteenth Amendment was ratified. However, the CSPI increased between 1919 and 1920 - the *latter* being the year when the Nineteenth Amendment was ratified. To mitigate this, we lag the amendment variable and the two instrumental variables by one year.

The diagnostic tests provide evidence that the amendment event variable in each model was properly instrumented. The weak instruments test statistic allows for the rejection of the null hypothesis that the instrumental variables are only weakly correlated with the endogenous regressor. The overidentification test assesses whether the instruments are “valid.” The statistics in both models fail to reject the joint null hypothesis in that the instruments are uncorrelated with the error term and that the excluded instruments are correctly excluded from the estimated equation.

The primary insight from these two models are the endogeneity test statistics. The results of these tests are only valid if the models are otherwise well specified, which is why the previous diagnostic tests are important. Here, the null hypothesis is that the specified endogenous regressors can be treated as exogenous, and the test statistics are too small to reject the null. Thus, there is no evidence that even the most significant constitutional amendments exert a significant effect on social capital levels.

5.3 Models

In the analysis that follows, we use the `brms` package (Bürkner 2017, 2018, 2021) in R (R Core Team 2021) to estimate the models. We use the `tidyverse` universe of packages for data manipulation (Wickham et al. 2019), the `posterior` package for summarizing posterior distributions (Bürkner et al. 2022; Vehtari et al. 2021) and the `ggplot2` package for plotting results (Wickham 2016). We discuss the specification and estimation of each of the three different models below.

5.3.1 Cross-national Models

Currently, the state of the art methodologically is Tsebelis (2022), where amendment rates are modeled with a heteroskedastic linear model to account for the changing residual variance as a function of independent variables. This strategy, while common in economics, has several disadvantages. The same problems that are evident in the linear probability model of binary variables also exist (though perhaps to a lesser degree) when modeling proportions (i.e., rates) as the dependent variable. While it would be possible to transform the probabilities into something unbounded using the logit transform (or similar) $y^* = \log(\frac{p}{1-p})$, in this case 20% of the observations are boundary values of 0 or 1 such that the logit transform is undefined: e.g., $\log(\frac{1}{1-1}) = \infty$. Again, we could solve this by shrinking the range (e.g., $p^* = 0.98 \times p + .01$) and then treat $y^* = \log(\frac{p^*}{1-p^*})$, but the amount by which we shrink the range (0.98 in the previous equation) is arbitrary and potentially partly determinative of the model result. While we *could* use this model in a pinch, there is no particularly good reason not to use a model that avoids making questionable assumptions.

We choose the negative binomial regression model (NBRM) as our starting point. Inherent in this model is the idea that the variance is related to the independent variables. Specifically, we use the following model:

$$\begin{aligned}
y_i &\sim NB(\mu_i, \phi) \\
&= \binom{n + \phi - 1}{n} \left(\frac{\mu_i}{\mu_i + \phi} \right)^{y_i} \left(\frac{\phi}{\mu_i + \phi} \right)^\phi \\
E(y_i) &= \mu_i \\
Var(y_i) &= \mu_i + \frac{\mu_i^2}{\phi}
\end{aligned}$$

In a conventional NBRM, we parameterize the expected value as $\mu_i = e^{\beta_0 + \beta_1 x_{i1} + \dots + \beta_k x_{ik}}$. The ϕ parameter (or more appropriately $\frac{1}{\phi}$) is the overdispersion parameter. In our Bayesian model we can also parameterize this term such that it is a function of other variables in a similar fashion to μ_i : $\phi_i = e^{\gamma_0 + \gamma_1 x_{i1} + \dots + \gamma_k x_{ik}}$. Doing so allows both terms that contribute to the residual variance (μ and ϕ) to be different functions of the variables of interest. We use this approach when warranted (not all models will benefit from this additional complexity).

In these models, we parameterize both μ_i and ϕ_i as follows:

$$\mu_i = \exp(\beta_0 + \beta_1 \text{Rigidity}_i + \beta_2 \text{Social Capital}_i + \beta_3 \log(\text{Words}_i) + \log(\text{Age}))$$

$$\phi_i = \exp(\gamma_0 + \gamma_1 \text{Rigidity}_i + \gamma_2 \text{Social Capital}_i)$$

Since the conditional variance of y_i is a function of both μ_i and ϕ_i , we can calculate how the variance relates to the variables of interest - Constitutional Rigidity and Social Capital.

Considering the model specification, note the inclusion of $\log(\text{Age})$. This is the exposure or offset term in the model. In the data, the amendment rate is calculated by dividing the number of years in which amendments were ratified by the number of years since the adoption of the current constitution. Including an exposure term allows us to generate a model that predicts rates instead of counts, like Tsebelis (2022), but without needing to make assumptions about the errors. This is particularly useful given the considerable variation in the age of constitutions.

We use an average first difference approach, changing a variable of interest while holding all other variables at their observed values (Hanmer and Kalkan 2013). The general approach works as follows: if we are interested in the effect of social capital, we would generate a new data frame (D_1) where for all observations we subtract $.5 \times sd(\text{Social Capital})$. We then generate a second new dataframe (D_2) where for each observation we increase social capital by the same amount ($.5 \times sd(\text{Social Capital})$). We then produce the relevant design matrices for each data frame: X_1 and X_2 , respectively. We take the $s \times k$ matrix of posterior draws of the model parameters ($\tilde{\beta}$) and generate predictions:

$$\begin{aligned}
\tilde{\mu}_1 &= e^{X_1 \tilde{\beta}'} \\
\tilde{\mu}_2 &= e^{X_2 \tilde{\beta}'}
\end{aligned}$$

We then take the difference between the two $\Delta = \tilde{\mu}_2 - \tilde{\mu}_1$. The Δ matrix is now $n \times s$ where n is the number of observations in the data and s is the number of posterior draws from the model parameters. Taking the average in each column gives a vector of length s of posterior average first differences in predictions. We can summarize this vector (i.e.,

calculate the average, quantiles, and posterior probabilities) to learn about the expected effect of the variables of interest. The procedure described above is how we would proceed to calculate the effect for the mean. We use a similar procedure to calculate the effect for the variance.

Table A4: Negative Binomial Models of Amendment Ratification

Predictor	Government Confidence	Party Confidence	Court Confidence	Group Member.	Civic Activism	Pol. Trust Index
Const. Rigidity	-1.72*	-1.89*	-1.50*	-1.45*	-1.52*	-1.76*
	(-2.40, -1.01)	(-2.61, -1.13)	(-2.19, -0.78)	(-2.24, -0.62)	(-2.24, -0.78)	(-2.46, -1.06)
Social Capital	0.71*	0.60	0.36	0.14	1.29*	0.20*
	(-0.04, 1.51)	(-0.40, 1.64)	(-0.23, 0.99)	(-0.12, 0.40)	(0.80, 1.78)	(-0.04, 0.45)
log(# Words)	0.55*	0.69*	0.55*	0.62*	1.02*	0.61*
	(-0.02, 1.12)	(0.12, 1.29)	(-0.02, 1.13)	(0.03, 1.23)	(0.50, 1.56)	(0.03, 1.19)
Intercept	-3.81*	-3.81*	-3.29*	-2.98*	-5.18*	-2.38*
	(-6.75, -0.94)	(-7.27, -0.57)	(-6.18, -0.58)	(-5.59, -0.44)	(-7.62, -2.85)	(-4.87, 0.04)
Shape: Intercept.	7.31*	8.01*	7.66*	2.45*	1.72*	0.88
	(2.34, 12.53)	(2.63, 13.82)	(2.77, 13.44)	(0.25, 4.83)	(-0.42, 3.92)	(-0.81, 2.54)
Shape: Const. Rigid	-0.17	0.13	0.02	-0.66	-1.27	-0.00
	(-2.04, 1.70)	(-1.83, 2.10)	(-1.87, 2.05)	(-2.71, 1.47)	(-3.44, 0.90)	(-1.81, 1.86)
Shape: Social Capital	-2.77*	-3.72*	-2.76*	-0.73*	0.71	-0.99*
	(-4.85, -0.76)	(-6.44, -1.12)	(-5.04, -0.82)	(-1.34, -0.17)	(-0.86, 2.20)	(-1.67, -0.34)
LOO IC (SE)	342.1 (18.3)	344.3 (19.7)	343.9 (20.2)	345.9 (19.2)	331.8 (18.7)	341.4 (18.9)
$R_{y,\hat{y}}$	0.82	0.80	0.85	0.87	0.90	0.82

Notes:

* indicates posterior probability greater than 0.9 on the same side of zero.

Dependent variable in all models is the number of constitutional amendments ratified.

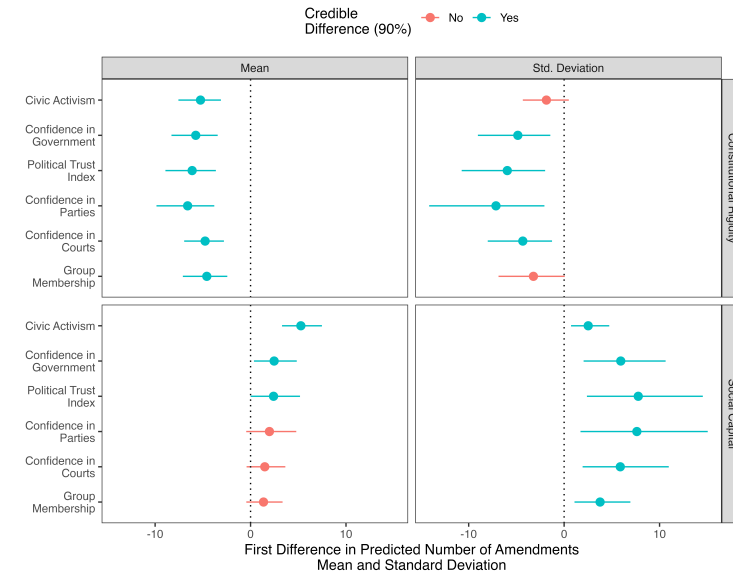
Number of observations in each model is 57.

LOO IC is an approximation to leave-one-out cross-validation - a comparative measure of model fit (akin to the AIC or DIC).

$R_{y,\hat{y}}$ is the correlation between observed and posterior mean fitted values.

Figure A4, below, shows the average first differences for each of the variables of interest for both the mean and variance. Observations are marked as “credible” if they have at least 90% of their posterior density on the same side of zero. The idea of a two-tailed p-value is foreign to the Bayesian paradigm, so we calculate the proportion of the posterior distribution that lies on the same side of zero as the posterior mean. We refer to this as the Bayesian p-value or the posterior probability. We calculate 80% credible intervals that allow visual tests against zero at the 90% Bayesian p-value level. In the interest of parsimony, we present a subset of these models in the paper - only those on the political trust index, group activity and civic activism.

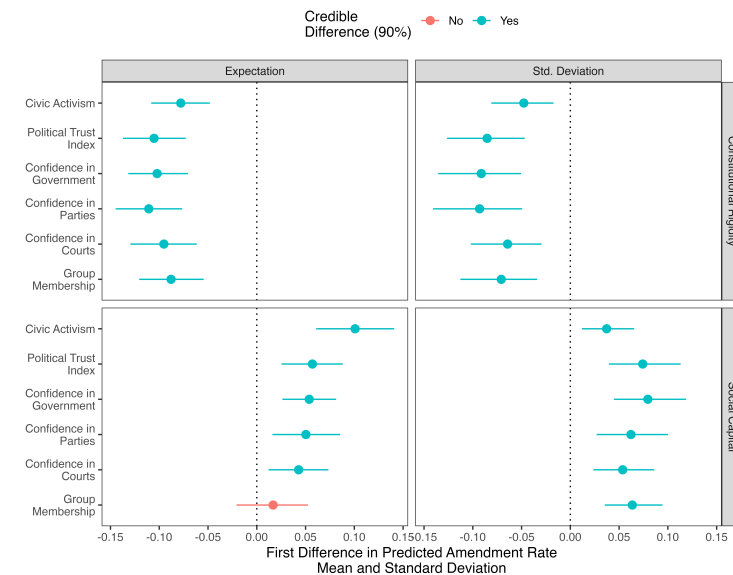
Figure A4: Effects of Social Capital on Mean and Residual Variance



5.3.2 Replicating the Heteroskedastic Linear Model

We also produce the same results as above for a heteroskedastic linear model, like the one Tsebelis (2022) used. The main difference here is that we do so using Bayesian simulation rather than maximum likelihood estimation.

Figure A5: Effects of Social Capital on Mean and Residual Variance from Heteroskedastic Linear Models



The heteroskedastic linear model produces quite similar results. In general, everything

Table A5: Heteroskedastic Linear Models of Amendment Ratification - Cross-national Data

Predictor	Government Confidence	Party Confidence	Court Confidence	Group Member.	Civic Activism	Pol. Trust Index
Const. Rigidity	-0.36*	-0.40*	-0.34*	-0.32*	-0.28*	-0.38*
	(-0.51, -0.22)	(-0.55, -0.24)	(-0.50, -0.19)	(-0.47, -0.17)	(-0.42, -0.14)	(-0.53, -0.23)
Social Capital	0.19*	0.21*	0.13*	0.02	0.31*	0.06*
	(0.06, 0.33)	(0.01, 0.40)	(0.00, 0.25)	(-0.04, 0.08)	(0.15, 0.48)	(0.02, 0.11)
log(# Words)	0.07	0.09	0.09	0.14*	0.20*	0.08
	(-0.06, 0.19)	(-0.04, 0.23)	(-0.05, 0.24)	(-0.01, 0.28)	(0.07, 0.34)	(-0.05, 0.21)
Intercept	-0.13	-0.18	-0.15	-0.08	-0.52*	0.27
	(-0.77, 0.51)	(-0.91, 0.53)	(-0.91, 0.61)	(-0.72, 0.54)	(-1.16, 0.10)	(-0.34, 0.86)
Sigma: Intercept.	-3.21*	-2.48*	-2.43*	-1.14*	-1.22*	-0.42
	(-4.67, -1.67)	(-3.86, -1.02)	(-3.84, -0.97)	(-1.81, -0.44)	(-1.89, -0.51)	(-1.03, 0.19)
Sigma: Const. Rigid	-1.39*	-1.37*	-1.02*	-1.12*	-0.86*	-1.33*
	(-2.10, -0.67)	(-2.11, -0.64)	(-1.70, -0.34)	(-1.80, -0.43)	(-1.59, -0.14)	(-1.99, -0.66)
Sigma: Social Capital	1.24*	1.09*	0.72*	0.36*	0.59*	0.34*
	(0.57, 1.88)	(0.36, 1.83)	(0.21, 1.24)	(0.16, 0.56)	(0.08, 1.11)	(0.16, 0.53)
LOO IC (SE)	-17.2 (14.5)	-7.7 (17.5)	-6.6 (17.3)	-11.6 (14.4)	-22.4 (13.5)	-13.7 (15.7)
$R_{y,\hat{y}}$	0.35	0.33	0.37	0.37	0.62	0.35

Notes:

* indicates posterior probability greater than 0.9 on the same side of zero.

Dependent variable in all models is the number of constitutional amendments ratified.

Number of observations in each model is 57.

LOO IC is an approximation to leave-one-out cross-validation - a comparative measure of model fit (akin to the AIC or DIC).

$R_{y,\hat{y}}$ is the correlation between observed and posterior mean fitted values.

that produces credible differences in the negative binomial model does so here. However, there are some instances when the converse is not true. Both confidence in parties and confidence in courts reliably increase the expected rate of amendments. Likewise, group activity and civic activism tend to reliably decrease the residual variation here where they did not in the negative binomial models. We find these two sets of results largely mutually corroborative. To the extent that there are differences, we feel the NBRM should be preferred as it makes a more tenable set of assumptions about the data generating process.

5.3.3 Cross-national Longitudinal Models

In the cross-national, longitudinal models, the dependent variable is binary, indicating country-years in which *at least* one amendment was ratified. We employ a dichotomous dependent variable because other studies (Ginsburg and Melton 2015; Tsebelis 2022) assume that multiple amendments are combined into a single legislative package. This is the same raw data that motivated our previous cross-national analysis. The main difference is that in this model, we have a time-varying measure of social capital from the Varieties of Democracy (V-Dem) project (Coppedge et al. 2021). Since the dependent variable is binary, we use as a starting point the model described in Beck, Katz, and Tucker (1999). They suggest using a spline or spell dummies build a flexible model of the time dependence. Carter and Signorio (2010) suggest that for most applications a third-degree polynomial is sufficiently flexible, particularly when coupled with the lack of familiarity most quantitative scholars have with splines. Our model is well-suited to treating the spells as a random effect - this provides incredible flexibility in terms of functional form, but the shrinkage inherent in these models will keep sparsely populated spell values closer to the global mean of the

random effect. Thus, it is a nice compromise between the (semi-)parametric model and the spell dummy variables leveraging the most beneficial features of each.

In the Beck, Katz and Tucker (1999) model, only the subset of the data where an $y_{i,t-1} = 0$ is used. This operationalizes the survival model that is being replicated in the logistic regression setup. As a reminder, a survival model tries to estimate the hazard - the probability that an event happens at time t conditional on it not having already happened. This makes the most sense in the health settings in which this model was developed, where the event is death or illness. Presumably if you're dead at time t you will continue to be dead at time $t + 1$. Illness is a bit trickier, but presumably if you are ill at time t and also ill at time $t + 1$, we could consider this time-frame as belonging to the same illness. With the case of amendments, the logic is even less clear. If an amendment is ratified at time t and at time $t + 1$, we know for certain those are *different* amendments. It is possible to consider amendment spells as times where a society is particularly prone to ratify amendments - enough so that it continues to ratify amendments each year for several years in a row. This may be a theoretical curiosity, though empirically it is difficult to test.

In addition to modeling the spell variable with a random effect, we also adopt the advice (at least in part) of Bell and Jones (2015). They suggest using both within- and between-transformations of the variables of interest. This permits the disentanglement of the between- and within- effects of those variables generating a more nuanced set of findings. The between-transformation of a variable X with observations i in groups j is just:

$$\bar{x}_j = \frac{1}{n_j} \sum_{i=1}^{n_j} x_{ij}$$

The within-transformation simply subtracts the between transformation from the observations value, x_{ij} . We will denote this as follows:

$$x_{ij}^{(w)} = x_{ij} - \bar{x}_j.$$

We can then specify the model as:

$$\log \left(\frac{Pr(y_{it} = 1)}{1 - Pr(y_{it} = 1)} \right) = \beta_{0i} + \beta_1 \text{Social Capital}_{it}^{(w)} + \beta_2 \text{Const. Rigidity}_{it}^{(w)} \times \text{Time-Varying}_i + \theta_s$$

$$\beta_{0i} = \gamma_{00} + \gamma_{01} \text{Time Varying}_i + \nu_{i0}$$

where θ_s is the random effect for the spell variable. In this iteration of the model, constitutional rigidity has within-variation for 16 countries. We use an interaction with a binary indicator identifying countries that have time-varying constitutional rigidity measures. We do so in a way that is often ill-advised (Berry, Golder, and Milton 2012), though in this particular application it makes sense. The conventional advice is to include all constitutive terms in the interaction (which would have us include the non time-varying values of constitutional rigidity as well). However, in this case since there is no time-varying information to be leveraged, the effect of the within-transformed rigidity variable for the countries without time-varying information is (and must be) precisely zero. Our model captures this.

In Model 1, we estimate the effect for all countries in the data, with an interaction for

those with time-varying information on constitutional rigidity. In Model 2, we subset the data to include *only* those countries with time-varying information on constitutional rigidity. This is meant as a robustness check for the effect of social capital.

In Models 3 and 4, we replicate Models 1 and 2, but include several of the control variables used by Tarabar and Young (2021) - namely, executive constraints, the log of GDP/capita, GDP growth, conflict and ethnic fractionalization. We subject these variables to a within-country transformation. The random intercept will pick up any important between-country variation that may be partly explained by the country averages of these variables. For two countries, India and Slovenia, we do not have ethnic fractionalization measures, so instead of removing them from the analysis, we simply set their within-transformed ethnic fractionalization measures to zero. This would indicate a constant ethnic fractionalization rate within country, though it is silent on the level at which that rate is constant.

We find that this effect is of roughly the same size. The only important difference exists between Models 2 and 4, where the effect of Social Capital becomes nearly zero when including controls for the countries with time-varying constitutional rigidity measures. Given that this is a small subset of countries in the main analysis, we do not see this as particularly troublesome, as the main finding holds moving from Model 1 to Model 3.

Nominally, we are estimating the probability of an amendment being ratified at time t , but we can also predict the expected amount of time it would take for an amendment to get ratified under different scenarios. Smaller intervals until amendments are ratified means a greater number of amendments ratified over a country's lifespan. We use a similar idea to the one described above to calculate the effect of variables in the model. This model is more complex than the previous one and as such permits the comparison of two quantities of interest.

5.3.4 Change in Probability of Ratification

First, we can calculate the change in the probability of ratification for each country year and average those changes across all years of the country's democratic history since the enactment of its current constitution.

Of the 80 countries included in the data, 78 have changes in predicted probability (in the aggregate) that have posterior probability greater than 0.9. The differences here range from

Table A6: Estimates of Cross-national Longitudinal Binary TSCS Models

Term	Model 1	Model 2	Model 3	Model 4
Intercept	-1.43* (-3.67, 0.79)	-7.77* (-13.82, -1.67)	-1.28 (-3.65, 1.03)	-7.92* (-14.63, -1.33)
Soc. Capital (within)	4.24* (2.47, 6.06)	4.57* (1.69, 7.76)	3.49* (0.61, 6.55)	-0.28 (-7.26, 6.98)
Const. Rigidity (within)	-3.46* (-7.49, 0.23)	-3.01* (-6.63, 0.46)	-3.77* (-7.95, -0.08)	-3.65* (-7.62, 0.15)
Const. Rigidity (between)	-2.81* (-4.53, -1.25)	-1.29* (-2.95, 0.09)	-3.15* (-4.97, -1.47)	-2.14* (-4.20, -0.45)
log(# Words)	0.19* (-0.01, 0.41)	0.72* (0.16, 1.31)	0.20* (-0.01, 0.42)	0.80* (0.20, 1.44)
Time Varying	0.18 (-0.61, 0.99)		0.21 (-0.65, 1.14)	
Exec. Const. (within)			-0.29 (-0.73, 0.18)	-0.43 (-1.74, 0.84)
log(GDP/capita) (within)			0.35* (-0.08, 0.78)	1.16* (-0.01, 2.33)
GDP Growth (within)			3.09* (-0.11, 6.34)	3.95* (-2.09, 10.24)
Conflict (within)			0.70* (-0.28, 1.68)	0.59 (-0.95, 2.15)
Ethnic Frac. (within)			-1.94 (-6.00, 1.94)	-4.14* (-10.61, 1.77)
sd(Country RE)	1.23* (0.94, 1.60)	0.45* (0.02, 1.12)	1.32* (1.00, 1.72)	0.47* (0.03, 1.16)
sd(Spell RE)	0.11* (0.00, 0.35)	0.22* (0.01, 0.67)	0.11* (0.00, 0.35)	0.21* (0.01, 0.66)
N	2002	569	1529	397
N Countries	80	16	78	16
LOO IC (SE)	1659.6 (52.0)	445 (30.1)	1327.5 (44.7)	385 (25.8)
PRE	0.12	0.06	0.12	0.10

Notes:

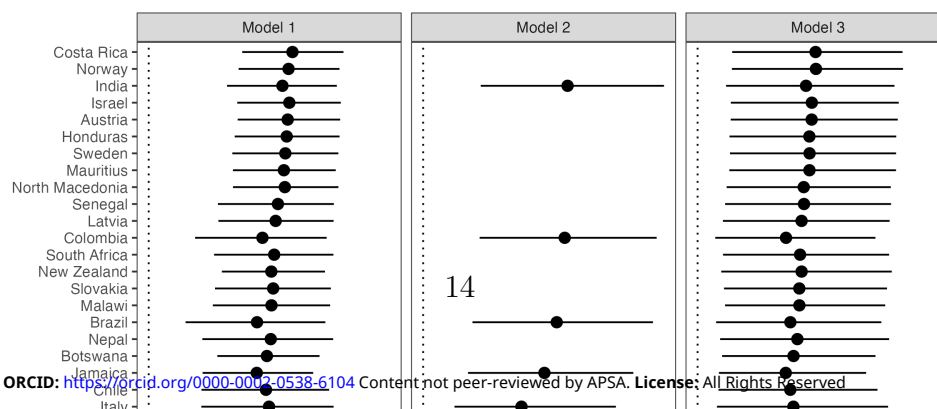
* indicates posterior probability greater than 0.9 on the same side of zero.

Dependent variable in all models is whether a constitutional amendment happened in each year (0/1).

LOO IC is an approximation to leave-one-out cross-validation - a comparative measure of model fit (akin to the AIC or DIC).

PRE is the proportional reduction in error

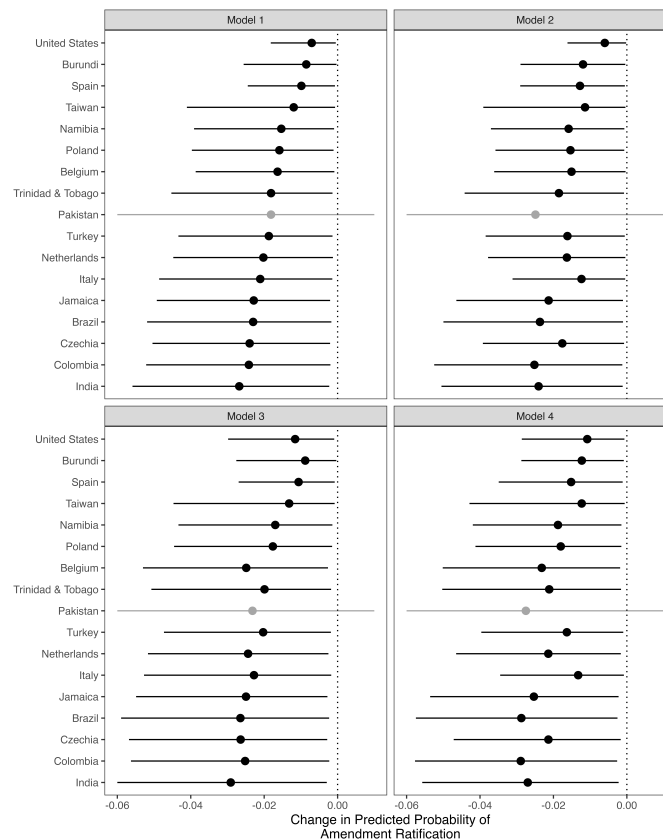
Figure A6: Effects of Social Capital on the Probability of Ratification



below 0.01 (e.g., Denmark, Niger and South Korea) to more than 0.08 (e.g., Israel, Austria, Honduras and Costa Rica). We could also look at this in the second model, using only the observations with time-varying values of constitutional rigidity. Figure A6 identifies those countries and the change in probabilities from Models 1, 2 and 3. The findings across the three models are quite similar. There are only two countries (France and Luxembourg) that had credible effects in Model 1 but not Model 3. Further, the ordering of effects is quite similar, which would suggest no major differences in the random effects across the two models. This is further corroborated by the similar standard deviations of the fixed effects across corresponding models with and without controls.

We could do the same thing for the constitutional rigidity measure. Using the within-country measure, each of the 16 countries included in the data have changes in predicted probability (in the aggregate) that have posterior probability greater than 0.9. On average these credible changes in predicted probability range from United States with a change of -0.01 to India with a change of -0.03. We also do this for Model 2, with only time-varying observations on constitutional rigidity and find that the effects are quite similar. Figure A7 also shows the same results for Models 3 and 4 which include control variables.

Figure A7: Effects of Constitutional Rigidity on the Probability of Ratification



The results across the four models displayed in Figure A7 are quite similar. In each panel, the results are ordered by the findings from Model 1. To the extent that the results change ordering, that indicates a change in the random effect - resulting in larger or smaller effects

given the same data and similar sized coefficients. We see this is largely corroborating the importance of constitutional rigidity across models.

We also estimated models that parallel Models 1-4, but include a referendum dummy variable and an interaction with social capital. We find no strong evidence of an interaction effect. Table A7 shows the results of these models.

Table A7: Models with Referendum-Social Capital Interaction

Term	Model 5	Model 6
Intercept	-5.43*	-5.71*
	(-10.16, -1.47)	(-10.98, -1.05)
Const. Rigidity (within)	-3.52*	-3.88*
	(-7.73, 0.27)	(-8.18, -0.03)
Soc. Capital (within)	3.38*	5.12*
	(1.22, 5.68)	(1.07, 9.40)
Referendum	-0.07	-0.36
	(-0.60, 0.50)	(-1.01, 0.26)
Soc. Capital (within):Referendum	-0.62	-2.36
	(-5.26, 3.87)	(-8.47, 3.52)
Const. Rigidity (between)	-2.62*	-2.81*
	(-4.31, -0.98)	(-4.63, -1.05)
log(# Words)	0.60*	0.66*
	(0.21, 1.06)	(0.22, 1.18)
Time Varying	-0.03	-0.08
	(-0.83, 0.78)	(-0.96, 0.73)
Exec. Const. (within)		-0.20
		(-0.75, 0.36)
log(GDP/capita) (within)		0.29*
		(-0.14, 0.72)
GDP Growth (within)		3.20*
		(-0.07, 6.69)
Ethnic Frac. (within)		-2.20
		(-6.37, 1.79)
Conflict (within)		1.02*
		(-0.09, 2.13)
sd(Country RE)	1.18*	1.24*
	(0.88, 1.55)	(0.92, 1.63)
sd(Spell RE)	0.11*	0.12*
	(0.00, 0.33)	(0.00, 0.40)
N	2002	1529
N Countries	80	78
LOO IC (SE)	1570.3 (50.8)	1259.9 (44.3)
PRE	0.13	0.15

Notes:

* indicates posterior probability greater than 0.9 on the same side of zero.

Dependent variable in all models is whether a constitutional amendment happened in each year (0/1).

LOO IC is an approximation to leave-one-out cross-validation - a comparative measure of model fit (akin to the AIC or DIC).

PRE is the proportional reduction in error

The first indication of additive effects is that the interaction term is not statistically significant in either model, though it is in the proposed negative direction). Following Berry et. al. (2010), we calculate the second difference in probabilities with respect to the social capital and referendum. In neither model do we find that the average second difference is credible at the 90% level.

In the cross-national analysis, we added in cultural controls because it was plausible that controlling for those variables could change our results. In some cases they did and in other cases they did not. The same cannot be said of our cross-national longitudinal analysis.

Here, we explicitly use a within-country estimator to estimate the effect of social capital. As such, any country-level differences are eliminated by design, including time-invariant cultural factors. Including the cultural variables in these models could reduce the country-level random effect variance, but would not change at all the within country effects estimated in the model. Since we are not interested in the between-country effects, we choose not to burden the reader with another set of results that cannot be meaningfully different from the ones we presented above.

5.3.5 State Amendment Models

The final analysis uses amendment proposal and ratification rates in the US states. The benefit here is that many aspects that varied in the cross-national setting will naturally be held constant in this analysis. That said, there is still considerable variation in contexts across states. The analysis here comprises two separate models - one of amendment proposal and one of amendment ratification. In the proposal stage, a negative binomial model of event counts is estimated. In this model, there is no exposure or offset term because theoretically the number of amendments proposed is unlimited, at least in most states. Arkansas, Colorado, Illinois, Kansas, and Kentucky all have some limits on the number of amendments that can be placed on ballots. One way to deal with this is to treat the limited states as having exposure terms equal to the limit and those without limits as having very high exposure values (e.g., 10,000). This, however does not really capture reality. Even in states without limits, the time, energy and political capital required to move amendments to the ballot suggest that there is some effective limit that is well below ∞ or even 10,000. So, we calculate the model with all states without an offset. We also estimated a model only on those states without ballot limits. The results are substantively similar, so the ballot limits are not driving our findings in a meaningful way.

As above, we calculate within-state variables and between-state variables for both constitutional rigidity and social capital. We started with the same specification as above for the amendment ratification model. We estimated random slopes for social capital but found that the residual standard deviations were very small. We retain the random intercept for both the expectation and overdispersion terms. Further, at the proposal stage, there is very little within-state variation for constitutional rigidity. The variation is small enough that the modeled is rendered unable to generate meaningful estimates of the variable's effect. As such, we move to including the between-state effect to capture the relevant differences in constitutional rigidity. Specifically, we estimate the model below for amendment proposals:

$$\begin{aligned}
 y_{ij} &\sim NB(\mu_{ij}, \phi_{ij}) \\
 \log(\mu_{ij}) &= \beta_{0j} + \beta_1 VP_{ij}^{(w)} + \beta_2 SC_{ij}^{(w)} \\
 b_{0j} &= \gamma_{00} + \gamma_{01} VP_j^{(b)} + \gamma_{02} SC_j^{(b)} + \gamma_{03} \log(\text{Words}_j) + \nu_{0j} \\
 \log(\phi_{ij}) &= \theta_{0j} + \theta_1 SC_{ij}^{(w)} \\
 \theta_{0j} &= \delta_{00} + \epsilon_j
 \end{aligned}$$

where μ is the expected value and $\alpha = \frac{1}{\phi}$ is the overdispersion parameter. Both parameters are modeled as a function of the independent variables of interest.

For the ratification stage, we estimated the following model:

$$\begin{aligned}
 y_{ij} &\sim NB(\mu_{ij}, \phi_{ij}) \\
 \log(\mu_{ij}) &= \beta_{0j} + \beta_2 SC_{ij}^{(w)} + \log(\# \text{Proposed Amendments}) \\
 b_{0j} &= \gamma_{00} + \gamma_{01} VP_j^{(b)} + \gamma_{02} SC_j^{(b)} + \gamma_{03} \log(\text{Words}_j) + \nu_{0j} \\
 \log(\phi_{ij}) &\sim N(\delta_{00}, \sigma^2)
 \end{aligned}$$

After some model diagnostic checking, we arrived at the specification described above. We remove the within-effect of constitutional rigidity on amendment ratifications because it introduces a considerable amount of noise due to lack of variation and does not change the findings otherwise (only Florida changed its amendment rules during the timeframe covered in this dataset). We also removed the parameterization of the overdispersion parameter as none of the variables had a credible effect on the overdispersion. This suggests that the residual variance is well-captured by the parameterized mean and single overdispersion parameter.

Table A8: Results for Proposal and Ratification Models

Term	Proposal All Obs	Ratification	Proposal No Limits	Ratification
Intercept	-9.34* (-13.03, -5.63)	2.27* (0.05, 4.69)	-9.87* (-13.65, -6.17)	2.90* (0.39, 5.80)
Const. Rigidity (within)	-0.36 (-2.01, 1.26)		0.16 (-1.62, 1.89)	
Social Capital (within)	0.10* (0.01, 0.19)	0.03 (-0.05, 0.11)	0.07* (-0.02, 0.17)	0.03 (-0.05, 0.11)
Const. Rigidity (between)	0.36 (-1.50, 2.27)	-4.43* (-9.15, -0.23)	0.47 (-1.41, 2.47)	-5.57* (-11.27, -0.80)
Social Capital (between)	-0.01 (-0.24, 0.23)	-0.10* (-0.18, -0.01)	-0.01 (-0.26, 0.24)	-0.11* (-0.19, -0.02)
log(# Words)	0.96* (0.68, 1.23)	-0.03 (-0.11, 0.04)	1.01* (0.73, 1.30)	-0.04 (-0.11, 0.03)
Intercept (σ)	2.08* (1.45, 2.92)		1.99* (1.34, 2.89)	
Social Capital (within, σ)	0.72* (0.12, 1.40)		0.72* (0.14, 1.38)	
sd(State RE)	0.56* (0.42, 0.74)	0.04* (0.00, 0.11)	0.54* (0.39, 0.73)	0.04* (0.00, 0.12)
sd(State RE [Shape])	1.36* (0.75, 2.21)		1.37* (0.73, 2.27)	
N	576	452	516	405
N States	48	47	43	42
LOO IC (SE)	2367.9 (50.9)	1463.8 (26.5)	2169.4 (48.0)	1340.1 (25.9)
$R_{y,\hat{y}}$	0.88	0.97	0.88	0.97

Notes:

* indicates posterior probability greater than 0.9 on the same side of zero.

All models are negative binomial models with state random effects in both the mean and variance equations. Dependent variable in “Proposal” models is the number of proposed amendments.

Dependent variable in “Ratification” models is the number of ratified amendments.

LOO IC is an approximation to leave-one-out cross-validation - a comparative measure of model fit (akin to the AIC or DIC).

$R_{y,\hat{y}}$ is the correlation between observed and posterior mean fitted values.

These suggest that social capital has a small-ish effect in the proposal stage, but nothing in the amendment ratification stage. What we will see below is that at the proposal stage, the small effects can really accumulate over time to produce relatively large changes in the

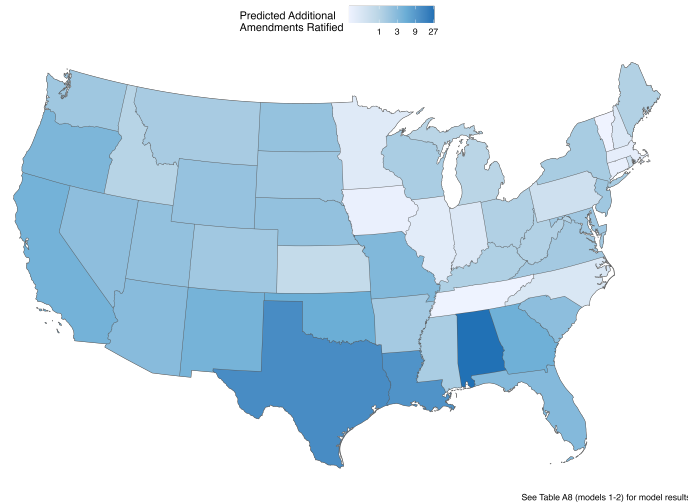
number of amendments proposed over the life of the constitution. In the simulation below, we first simulate the effect of a one standard deviation change in social capital in the same way as above. That produces two different sets of estimates - one under a lower social capital scenario and one under a higher social capital scenario. We then use the posterior means of those two predictions (rounded to the next highest integer) as the exposures in our predictions for Model 2. Again, we generate two sets of predictions - one under a lower level of social capital for everyone, using the lower social capital predictions from Model 1 as the exposure term, and one under a higher level of social capital for everyone, using the higher social capital predictions from Model 1 as the exposure term. This captures the total effect of social capital across the two models.

The figure below gives the change in the number of amendments ratified by state (aggregating across all years) for a one standard deviation change in within-state social capital. It is largely the Southern states that see a big increase in additional amendments with Alabama predicted to see more than 28 extra amendments proposed as a function of changes in social capital. Texas and Louisiana are also expected to experience double-digit increases (13 and 11, respectively). The changes here are due primarily to the increased number of amendments that are proposed over time. In this case, 32 of the 47 effects are statistically reliable. Specifically, predicted values bigger than 0.85 have posterior probabilities greater than 0.9.

One potential problem with the analysis above is that it does not incorporate the uncertainty in the first-stage predictions. We may imagine that this would be the appropriate thing to do, but there are arguments to the contrary as well. Since we built the first-stage model to test a hypothesis and not to produce a “good” model with the smallest posterior variability, there may be some unnecessary noise in those predictions owing to extraneous variables included. Nonetheless, we estimate the simulation this time, not using the posterior means but for each draw of the posterior for the predicted number of proposed amendments.

The figure below shows the result of this different simulation. Here, we get $\hat{P}_{i,t}^-$ and $\hat{P}_{i,t}^+$, which are the predicted proposal rates for each of the i states in each of the t years given a half standard deviation decrease in social capital (with a $-$ superscript) or a half standard deviation increase in social capital (with a $+$ superscript). We get one of these for each posterior draw from the MCMC algorithm. Next, we calculate $\hat{R}_{i,t}^-$ and $\hat{R}_{i,t}^+$, which is the predicted number of ratified amendments given a half standard deviation plus or minus in social capital *and* accounting for the change in proposed amendments. That is, $\hat{R}_{i,t}^-$ is calculated using $\hat{P}_{i,t}^-$ as the number of proposed amendments. The same is true for the $+$ superscript variables. This provides a much better accounting of the overall effect of social capital on ratifications through the proposal process. We find that increase in proposals translate into increases in predicted changes in ratifications. Alabama is expected to ratify roughly 28 more amendments over the years covered if social capital increases by a standard deviation. Of these 47 results, 22 are statistically reliable, with posterior probabilities greater than 0.9.

Figure A8: Effect of Social Capital on State Constitutional Amendment Ratifications
(alternative simulation)



5.4 Comparison with Tarabar and Young

One reviewer posed the question - what do we learn here that we did not already know from Tarabar and Young (2021)? To answer this question comprehensively, we must think about the nature of the social science research process in its entirety. We feel that our analysis is superior in several different ways.

5.4.1 Theory and Conceptualization

During the review process, there was a suggestion that our concept of social capital is really just a repackaging of Tarabar and Young's *amendment culture* concept. Theoretically, these are different ideas. Amendment culture, as defined by Ginsburg and Melton (2015) and used in Tarabar and Young (2021, 2) is "the set of shared attitudes about the desirability of amendment." We define *social capital* in this context as the "trust, reciprocity and civic activity produced by interpersonal networks [that] help elites, ordinary citizens, and social movements overcome the transactions costs" of ratifying amendments. These ideas are likely related, though culture is something that changes slowly over generations whereas social capital is more variable over time. If we are right, then higher levels of social capital *within a constitutional regime* should coincide with higher rates of constitutional amendments. Since culture is a much stickier concept, it will not be able to explain a phenomenon that changes in relatively small time-scales. The culture argument would have us believe that all of that variation across time is simply idiosyncratic. We feel that an adequate explanation of amendment rates must do better than that.

5.4.2 Measurement

Both our social capital measures and Tarabar and Young's amendment culture variables are derived from surveys. To some degree, the quality of the survey determines the quality of the measurement. Our social capital measures in the cross-national analysis are derived from the WVS, one of the most rigorous and longest-running cross-national surveys in the world.

The Hofstede indicators were originally collected from surveys of employees of IBM subsidiaries in 40 countries from 1966-1973.³ Hofstede stresses that for cross-country comparison, samples should be "matched" meaning that the same kinds of respondents should be used in each country, and while the scores may not replicate (i.e., students may have different values than grandparents), the inter-country differences can be reliably estimated. This strains credibility a bit. We could imagine that, on average, students or employees of a huge multinational corporation may exhibit similar patterns across countries, but would that hold for all groups? What about farmers, for example? If we could survey farmers in each of these countries, would these same cultural patterns obtain? Would we expect them to be present? Would not a nationally representative measure require a nationally representative sample? Assuming that students, corporate employees and farmers might respond differently, a nationally representative measure would depend on the relative prevalence of each group in society. Since societies have different proportions of students, corporate employees, farmers (and all other groups), we could not possibly expect these matched samples to produce something that is nationally representative.

Furthermore, for some unexplained reason, the Hofstede data are non-replicable. On his website, Geert Hofstede warns, "Some of the dimension scores obtained in replication studies fall outside the 0-100 continuum."⁴ Finally, scholars have criticized these Hofstede measures for presuming that national culture does not change over time, overestimating the number of cultural dimensions, and misinterpreting their meaning (Ailon 2008; Baskerville 2003; Baskerville-Morley 2005; Fang 2003; McSweeney 2002, 2009; Taras, Steel, and Kirkman 2012; Venaik and Brewer 2016).

On the whole, we feel that our survey-derived measures stand on firmer ground.

5.4.3 Modeling

So far, we have argued that both our conceptual framework that accounts for time-varying changes in amendment rates and our measures derived from the most rigorous cross-national survey project to date are superior to those used in Tarabar and Young (2021). The next question concerns the model. Tarabar and Young use an ordinary least squares (OLS) model to estimate amendment rates. We address the problems with using OLS on amendment rates elsewhere in the appendices, particularly with respect to Tsebelis' use of the heteroskedastic linear model. Those concerns, and others, are the same here. Consider model 8 in Table 4 (p. 12 of Tarabar and Young) and model 6 in Table 5 (p. 13 of Tarabar and Young). These models have 89 and 72 observations representing 43 and 35 countries, respectively. These models have a very high ratio of estimated parameters to number of observations, or to put

³ See: <https://geerthofstede.com/research-and-vsm/vsm-2013/>.

⁴ See <https://geerthofstede.com/research-and-vsm/dimension-data-matrix/>.

it differently, very few residual degrees of freedom. Model 8 (Table 4) estimates an intercept, 14 coefficients and six region fixed-effects for a total of 21 parameters. This is a lot to ask of 89 data points, especially since some of the measures will not vary within country, i.e., across constitutional regimes. Model 6 (Table 5) is even worse, estimating the 21 parameters in Model 8 plus six, for a total of 27 parameters on 72 observations - fewer than 3 observations per parameter. In a simple simulation, you can show that if you randomly generate X with 27 variables and $n = 72$, and then randomly generate a y that has no relationship to those 27 variables, a regression of y on X will yield an R^2 of, on average around .4 with, on average, one or two variables being significant and can be as high as around 0.7 with as many as 11 or 12 variables being significant.

We use a Bayesian negative binomial model that estimates seven parameters on 57 observations, using roughly 8 observations per parameter, which seems considerably more reasonable. We extol the virtues of that model elsewhere in the appendices. However, if we had done the same kind of simulation described above with 57 observations and 7 parameters, it would produce an OLS model with an average R^2 of around 0.1 with, on average, a single significant variable.

Comparing model performance is difficult, as the question cannot be answered in the conventional way - a kitchen-sink model containing both sets of variables. This would simply result in models that we feel would use too many degrees of freedom. There is not enough Hofstede data available to make statistically-reliable comparisons between the two approaches. Only 31 democracies in Tarabar and Young's sample also exist in ours, and only 24 of these have data for all four Hofstede culture dimensions. Furthermore, the democracies analyzed by Tarabar and Young are not a random sub-sample of the democracies included in our sample. Several nations that are important cases in any study of democratic constitutionalism are missing from Tarabar and Young's (2021) sample, including: France, India, Israel, Japan, and the United States. Deriving insights from such a model could easily lead us in the wrong direction.

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