There Is Only One Question: Diffuse Institutional Confidence in a Polarized America

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Abstract

Americans have broadly experienced over a decade of extraordinary institutional stresses, which has raised questions about the underlying structure of their trusts and confidences in those institutions. In this paper, we explore the degree to which Americans express a unidimensional confidence that spans multiple institutions in a mixed ecology of governance. Using a three-wave national panel spanning 2018 to 2023 and an ordered-categorical bifactor model, we study confidence in nineteen institutions across public and private authority. During this time, confidence increasingly behaves as a single dimension. A general factor G spans both public and private sectors and comes to dominate the confidence landscape, with the share of common variance it explains rising from 0.20 in 2018 to 0.47 in 2023. This consolidation occurs due both to institutions binding more tightly to G, and the sector-specific structure that once separated public from private evaluations receding. Establishing measurement comparability across parties and waves, we find that G became not only more dominant but more partisan. The gap between Democrats and Republicans roughly doubled, from 0.38 to 0.79 standard deviations, even as overall confidence declined only modestly (about 0.27 SD). This pattern is most consistent with a mechanism based on partisan sorting. Residual, item-specific politicization concentrates in a stable set of contested institutions and, for the executive and the military, reverses with control of government. Finally, G has important downstream consequences, as respondents who began with higher confidence grew more committed to democratic norms five years later. This is true at nearly identical magnitudes within each party, suggesting G is not proxying for partisanship. Diffuse institutional legitimacy has become a single but party-stratified disposition, suggesting

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a more brittle foundation for democratic resilience than the shared reservoir of support it is often assumed to be.

Introduction

Public confidence in institutions is a cornerstone of democratic governance and policy capacity. When citizens believe that authorities and rule-makers are competent, fair, and acting in the public interest, they are more willing to comply with rules, accept adverse outcomes, and cooperate with collective projects (Easton 1975; Levi and Stoker 2000; Tyler 2006; Hetherington 2005; Devine 2024). When confidence erodes, policy capacity weakens and democratic accountability can become brittle: citizens attribute poor outcomes to corruption rather than error, treat losses as illegitimate rather than routine, and interpret procedural constraints as partisan weapons rather than neutral guardrails (Hetherington 1998; Norris 2011; Achen and Bartels 2016; Daniller 2016; Citrin and Stoker 2018). The fundamental importance of this topic is reflected in how often political scientists and public opinion experts ask individuals about their confidence in a wide range of institutions. What remains at question, however, and what we focus this paper on answering, is the degree to which an individual's reported "confidence" attaches to any single institution, as opposed to representing a specific instantiation of a more general trust that extends beyond any one institutional authority. Put bluntly: Do citizens evaluate institutions one at a time, or is there only one general "institutional confidence?"

The stakes of that question have grown as the objects of confidence have multiplied. Citizens experience authority in a mixed system that includes elected officials and public agencies, but also major firms, banks, universities, media organizations, and other private or quasi-private institutions that exercise governance-like power (Cutler et al. 1999; Hall and Biersteker 2002; Büthe and Mattli 2011; Culpepper 2011; Gillespie 2018; Gorwa 2019). These actors write and enforce rules, limit and sanction behavior, and shape access to information and opportunity. Social media platforms moderate speech and coordinate "private ordering" of public discourse (Klonick 2018; Balkin 2018; Gillespie 2018; Gorwa 2019); credit rating agencies and financial institutions condition state capacity and distributive politics (Sinclair 2005; Fourcade and Healy 2017); universities structure life chances and social hierarchies (Collins 1979). If legitimacy is partly the public's belief that rule-makers are entitled to

rule and will do so on acceptable terms, then institutional confidence is increasingly about legitimacy across *both* public and private authority, not only political institutions in the narrow sense. Identifying a common dimension of institutional confidence across both public and private authorities would suggest not merely a general scale of political legitimacy, but a neo-Eastonian measure of the public's diffuse acceptance of rule-making and rule-enforcing at a societal level (Easton 1975).

Even remaining within the traditional governing institutions allows us to speak to one of the central debates in the study of democratic legitimacy. Over the past decade, scholars have argued whether advanced democracies face a genuine "crisis" of institutional trust or merely institution- and period-specific fluctuations (Valgarðsson et al. 2025; van der Meer 2017; Zmerli and van der Meer 2017). However, this question, too, requires we know whether "trust" is one thing or many. Simultaneous declines across many institutions are observationally equivalent to a broad decline in a single factor, but they do not diagnose the same democratic issues. The former can reflect coincendental isolated disappointments, all the time hiding the latter's suggestion of eroded diffuse legitimacy. This analytical difficulty extends to comparisons across groups and times, especially as certain institutions become politicized or polarized. To understand any of the fundamental questions about institutional confidence, we must first back out what that concept even means.

In this paper, we analyze the American Institutional Confidence Poll (the "AIC Poll") to answer this question. A three-wave national panel that measured the same respondents' confidence in nineteen institutions spanning public and private authority in 2018, 2021, and 2023, the AIC Poll allows us to observe whether a general factor reaches across sectors rather than within them. Its panel design, spanning a five-year stretch of exceptional institutional stress, allows us to ask whether such a factor exists, whether it strengthens over time, for whom, and to what end. The structure allows us to distinguish between three plausible mechanisms for a common single dimension of confidence: 1.) a trait, across a population, whereby individuals either trust or withhold trust generally, 2.) a correlated shock where individuals are viewing institutions separately, but the institutions all suffer from a sustained run of failures, and 3.) partisan sorting where party identity becomes increasingly powerful,

and used as a lens through which an ever larger number of institutions are assessed (see, e.g. Hetherington and Rudolph 2015; Ladd 2012; Levendusky 2013; Lelkes 2016; Klein 2020).

Our analyses reveal that a single general factor is both present and increasingly dominant. In all three waves, we use a bifactor model that separates a general confidence factor (which we call "G" hereafter) from residual clustering within the public and private sectors. This model fits the nineteen items far better than a model built from sector-specific factors alone. Over time, the share of common confidence variance attributable to G rises from 0.20 in 2018 to 0.41 in 2021 to 0.47 in 2023. This growing dominance reflects a real strengthening of the general factor. In 2018, G is comparatively weak and institutions are only loosely tied to it; over the panel they come to load on it far more heavily, even as the sector-specific structure that once organized confidence recedes. A single dimension increasingly absorbs evaluations respondents once kept apart. Having established the existence and growing explanatory power of the single general factor, we then analyze the factor's shifts over time. We find that confidence declines over the period, but only modestly, by roughly 0.27 SD. The larger movement is distributional in nature: the gap in G between Democrats and Republicans, estimated on a common latent scale, roughly doubles, from 0.38 SD in 2018 to 0.79 SD in 2023. Combined, these findings suggest partisan sorting as the driving force behind the rise of uni-dimensionality.

Finally, we show one way in which G matters practically. Exploiting the panel, we find that respondents who began the period with higher G became, five years later, more trusting of government, more satisfied with democracy, and more willing to grant the opposing party legitimacy. These effects are nearly identical within each party (the effect on government trust is $b = .093$ among Democrats, .097 among Independents, and .085 among Republicans), so G is not merely a proxy for partisanship. Within individuals, five-year changes in G and in government trust move together at $r = .259$. Generalized confidence thus behaves like a reservoir of diffuse support: a resource for democratic commitment that operates within partisan camps, not only between them.

The remainder of the paper proceeds as follows. The next section situates the central question of whether Americans hold one institutional confidence or many within existing

scholarship on political trust, affective polarization, and Eastonian diffuse support. We then introduce the AIC Poll and our measurement strategies: the ordered-categorical bifactor model, the multi-group and over-time invariance tests that render partisan and temporal comparisons credible, the MIMIC screens that localize residual politicization, and the cross-lagged design behind the downstream analysis. The results section takes each analysis in turn, before we conclude by asking what a unified, polarized, and consequential dimension of confidence implies for legitimacy in a mixed ecology of public and private authority. We begin, though, by examining the theoretic underpinnings of institutional confidence more broadly.

Background and Theory

In this section, we lay the conceptual groundwork for assessing the dimensionality of institutional confidence. We start with what institutional confidence is, and why its structure is a question about both legitimacy and measurement. We then discuss why the objects of that confidence now reach beyond government to the private and quasi-private institutions that increasingly govern public life, and what could bind evaluations of such disparate institutions into one dimension. Finally, we consider why the answer matters for democracy. Is a generalized confidence in institutions merely a description of a polarized public or a resource on which citizens' commitment to democratic norms depends?

Institutional Confidence and Diffuse Support

Institutional confidence is a judgment of whether an institution that governs public life is competent, fair, and acting in the public interest, and thus whether it deserves the deference of those it governs (Hetherington 2005; Levi and Stoker 2000; Citrin and Stoker 2018). It is distinct from social trust, which concerns the trustworthiness of other members of the public, *and* from approval of particular leaders or policies, which rises and falls with day-to-day performance. What makes institutional confidence politically consequential is that

it attaches to the standing arrangements through which collective decisions are made and enforced, rather than the individual or individuals occupying them (Miller 1974; Norris 2011).

The competence judgment at the heart of reported institutional confidence can come from two different sources. First, the judgment stems in part from how the reporter views the "record" of the institution, i.e. how the institution has performed recently. Second, the level of reported confidence is colored by the general orientation of the reporter toward institutional authority more broadly. Easton calls the former specific support and the latter diffuse support (Easton 1965, 1975). Specific support is performance-contingent and bound to a particular object (here, an institution); diffuse support is a standing, slow-moving reservoir of legitimacy that survives poor performance and attaches to *systems* of objects (in this case, a group of institutions) (Levi and Stoker 2000; Tyler 2006). Most confidence judgments blend the two, so the same rating can register a verdict on what an institution has done or a general inclination to grant or withhold legitimacy.

In order to understand the major takeaways from a battery of institutional confidence questions asked to a set of respondents, one must therefore wrestle with a dimensionality question: to what degree is each respondent pulling from a diffuse support for institutionality, rather than remarking on the specific outcomes associated with an individual institution? If institutional confidence responses are mostly specific in nature, ratings are largely independent and no common reservoir exists; if diffuse support does much of the work, a single underlying orientation binds the ratings together. Identifying whether such a general orientation exists, and assessing whether it is growing, become questions not only about proper interpretation of the ratings, but about how the population views institutional legitimacy writ large in contemporary America. We review the literature on that topic and why it matters in the following subsection.

One Confidence, or Many?

Research on political trust has frequently stood on either side of the dimensionality issue without directly acknowledging it or adjudicating between the two options. Much research proceeds as if confidence were essentially one thing, aggregating items into a single index

of "trust in government" or reading a lone survey question as a summary of institutional legitimacy (Hetherington 2005; Citrin and Stoker 2018). Other work proceeds as if confidence in a particular institution should be tracked separately on the grounds that courts, legislatures, police forces, etc., all answer to different publics and perform on different terms (Norris 2011).[1] Each approach assumes an answer to the dimensionality question without testing one.

The stakes of these assumptions have become clearest recently in the debate over whether advanced democracies face a genuine crisis of institutional trust or "merely" a churn of institution- and period-specific fluctuations (Valgarðsson et al. 2025; van der Meer 2017; Devine and Valgarðsson 2024). The distinction is important, as the former signals an erosion of diffuse legitimacy and a great deal of difficulty in repair, while the latter suggests that fixes may be available at the institution-level, if necessary at all. However, trend data alone cannot tell the two stories apart, and credible arguments can be made for either side *using the same data*. This is especially true as the number of institutions producing disappointing results increases. Ultimately, without being able to separate the common component of institutional confidence apart from the institution-specific one, debate stalls on evidence consistent with either story (Dalton 2004; Zmerli and van der Meer 2017).

Dimensionality, then, is a quantity to be estimated rather than a modeling convenience to be assumed. Rather than average the confidence items into an index or examine them one at a time, we model the battery so that a general factor and institution-specific factors are estimated jointly, focusing on the share of variance each accounts for, both within and across waves of our data. Recent work has begun to test this structure head-on. For example, Mancini et al. (2025) examine trust across both institutional and interpersonal targets, finding that a correlated many-factor model actually fits best. The article makes a strong case that different types of trusts and confidences may share common wellsprings, even as the authors in that piece pool very different objects. We restrict our attention to institutions, and add a panel that lets us watch the structure change. Rather than settling dimensionality once, we ask whether a general factor over institutional confidence grows more dominant over

[1] Recent structural work in this vein separates trust in "order" institutions from trust in "representative" ones, even once cross-group comparability is established (Stals et al. 2024).

time, whether it divides by party, and what it predicts. Throughout, we consider the size of the estimated general factor as a measure of how much of institutional confidence is diffuse. The construction of that confidence battery is non-trivial, with many institutions arguably playing a role in the way individuals perceive the legitimacy of the larger system. We turn to those concerns now.

The Objects of Confidence: Public and Private Authority

Modern citizens encounter authority in a mixed system, one in which private and quasi-private organizations exercise power that looks very much like governance (Cutler et al. 1999; Hall and Biersteker 2002; Büthe and Mattli 2011). Platforms write and enforce the rules of public discourse, deciding what may be said and who may say it (Klonick 2018; Balkin 2018; Gillespie 2018; Gorwa 2019); banks and rating agencies set the terms of credit that shape economic successes (Sinclair 2005; **?**); universities allocate opportunity and confer or withhold social standing (Collins 1979). These institutions are part of daily life, even when they operate in the background. They make and enforce rules, impose sanctions, and distribute advantage, and their claim to do so legitimately is contested on much the same terms as the state's (Bernstein and Cashore 2007; Black 2008; Suzor 2018).

Scholars of public trust have recognized this similarity in practice, if not always in theory. When Americans are asked about their confidence in institutions, they have long been asked about far more than government. The General Social Survey's confidence battery, in place since the 1970s, mixes Congress, the executive, the courts, and the military with banks, major companies, the press, organized labor, medicine, education, and organized religion (Smith 2008); Lipset and Schneider built an entire account of the "confidence gap" out of public attitudes toward business, labor, and government together (Lipset and Schneider 1983). The theory has frequently lagged behind, typically interpreting these responses as "political" trust, as if the private items were noise or a separate matter altogether (Citrin and Stoker 2018; Norris 2011).

If diffuse support is a generalized orientation toward authority rather than a judgment about any one office, there is no principled reason to stop measuring that support at the

boundary of the state. Rather, the relevant reservoir is confidence in the institutions that order public life, private institutions included. Of course, one way to identify *whether* it makes sense to think about a single legitimacy orientation overarching both private and public institutions is by including both in a full battery and testing for a general factor directly. Were we to identify such a factor, it would provide evidence that generalized legitimacy spans a mixed ecology of authority, and raise a question about why that might be the case. We discuss potential mechanisms for that immediately below.

Three Possible Mechanisms Behind a General Factor

If there exists any form of general factor over institutional confidence, its strength, and perhaps its detectability, is likely to change over time. This raises a question as to what dynamics or mechanisms might control those movements. We identify three processes that could produce a single dimension spanning institutions, with each dynamic carrying very different implications for what the movements of a general factor might mean.

The first potential mechanism is dispositional. Individuals may simply differ in a stable, trait-like propensity to extend or withhold trust, such that some people report high confidence across the board while others go low, largely without consideration of any institution's conduct. A general factor structured in this way is real but politically unimportant. It would be an aggregation of temperaments or attitudes rather than a signal about the standing of institutions in the population's view. From a diagnostic perspective, there would be little reason to expect its importance to grow systematically over time or to fall along partisan lines.

The second mechanism is correlated shock. Institutions need not fail in isolation; a pandemic, a financial crisis, a contested election, or a sustained stretch of poor governance can weaken performance across many institutions at once. This would produce a common component in confidence that reflects shared disappointments among respondents, without requiring the consistent attitude at the individual-level that the dispositional mechanism rests on. Instead, this dynamic can even see the general factor grow as shocks accumulate. Diagnostically, what distinguishes it is symmetry in the way movement happens. If simul-

taneous failures drive the general factor, then movement should be broadly shared across subgroups of respondents, pulling, e.g., Democrats and Republicans in the same direction rather than apart.

The final mechanism involves partisan sorting. As affective polarization deepens, party identity increasingly operates as a lens through which partisans evaluate the world, and the reach of that lens extends well beyond electoral politics (Iyengar et al. 2019; Levendusky 2013; Lelkes 2016). We know that partisanship structures attitudes toward Congress and the president, but there is evidence that it now organizes evaluations of the press, universities, technology platforms, and even science and public health (Ladd 2012; Gauchat 2012; Klein 2020). As institutions are increasingly sorted into partisan camps, a respondent's confidence across the whole battery will be driven less by what each institution does than by whether it is seen to be on the "right team." That sorting process would compact otherwise unrelated evaluations into a single dimension, and produce a different diagnostic signature. Here, the general factor should grow more dominant as it grows more divided by party, since the same sorting that unifies each side's evaluations drives the two sides apart (Hetherington and Rudolph 2015; Achen and Bartels 2016).

The three accounts share the potential to produce a general factor, but diverge along two dimensions our panel can observe. A dispositional factor should be stable and politically symmetric; a correlated-shock factor should move together across party; only partisan sorting predicts a general factor that becomes both more dominant and more partisan over time. Before we discuss our expectations for the existence and genesis of the general factor, we consider two final, connected questions: does a general factor even matter, and if so, for what?

Confidence as a Democratic Resource

In Easton's account, the reservoir of diffuse support allows a political system to survive popular disappointment. Citizens continue to accept institutional authority, work within established rules, and treat bad outcomes as losses to be reversed by legitimate means rather than as evidence the system is rigged (Easton 1975). If a general factor of institutional

confidence captures that reservoir, then it should shape how people act toward the political order, and in particular how firmly they hold to democratic norms when those norms are tested.

This has become more difficult as polarization has deepened. A growing body of work finds that when partisanship is intense, citizens face a tradeoff between their commitment to democratic principles and their "team" loyalty, with many resolving it in favor of party (Svolik 2019; Graham and Svolik 2020). Partisans are willing to tolerate, excuse, or even reward norm violations by their own leaders that they condemn in the opposition. The commonly expressed worry in democratic studies is that this is the mechanism through which mass polarization erodes democratic guardrails. Citizens do not become open authoritarians, but simply become unwilling to punish their own side for breaking the rules (Levitsky and Ziblatt 2018).

Generalized institutional confidence could serve as a counterweight. An individual with a deep reservoir of diffuse support has reason to extend legitimacy to institutions, and to political opponents, even when doing so may be costly to their party, and may be more willing to sustain democratic commitments under partisan pressure. Modern accounts of democratic deconsolidation tend to ask whether citizens' attachment to democracy itself is weakening (Foa and Mounk 2016). We posit that generalized confidence may be part of what holds that attachment in place, a resource for democratic resilience.

Whether generalized confidence *actually* predicts commitment to democratic norms is an empirical question, and we set out in this paper to answer it. We observe the same individuals over five years, allowing us to ask whether those who begin with higher generalized confidence become more committed to democratic norms later on, rather than merely correlating the two at a single moment. We can also disentangle whether generalized confidence is distinct from partisanship by looking at the degree to which confidence predicts democratic commitment *within* parties, rather than only across them. Before we describe our data and analytic approach, we quickly identify the specific expectations that guide the analyses to follow.

Expectations

Our theory yields three expectations alongside a related general research question that, when combined, organize the analysis. First, and fundamental to the rest of the project, we expect the confidence battery to reveal a single general factor spanning public and private institutions, rather than a set of sector-bound evaluations. As individuals experience constraints on their behavior from an increasingly complicated and intertwined number of public and private institutions, we expect their assessment of those institutions to be correlated, even across sectors. This correlation should yield a general factor that orders confidence in a wide range of our surveyed institutions.

Second, we expect that single general factor's dominance to grow over the panel. As complexity in the system increases, and more semi-private and private institutions become implicated both in governance and partisan struggles, we would expect individuals to associate those institutions together more openly and assess them as such. We want to focus carefully on how the growth of the dominance manifests, leaving as an open question whether there becomes a tighter bond between institutions and G, or a collapse of the previous sector-specific structure, or both?

Third, and most diagnostic of the mechanism at play, we expect the general factor to become more partisan as it becomes more dominant, producing a widening gap between Democrats and Republicans rather than a symmetric shift. If the association of institutions with a partisan struggle is in part what has bound them together, it stands to reason that this will also ensure that actual levels of G will diverge across the parties as this relationship tightens.

Finally, as another open research question that stems from our investigations, we want to know the degree to which an individual's level of our generalized institutional confidence relates to their commitment to democratic norms. If generalized confidence is to act as a democratic resource in the Eastonian sense, we should expect these to be connected, but have no clear expectations as to how, or whether this connection would hold both across

and within parties. We now turn to the data and methods that allow us to assess each expectation, as well as this final research question, in turn.

Data and Methods

We conduct five analyses to assess our expectations. Analyses I and II establish whether a general confidence factor exists and how its dominance changes over time. Analyses III and IV ask whether that factor is partisan in nature, and specifically whether Democrats and Republicans can be compared on a common scale, how large the gap between them may be, and which institutions carry partisan signal beyond the general factor. Analysis V asks whether the factor matters, testing whether earlier confidence predicts later commitment to democratic norms. In order to conduct these analyses, we need a particular type of panel survey that measures institutional confidence across a wide range of public and private institutions. The American Institutional Confidence Poll is exactly that, and we describe it in the first subsection below, before moving on to the broader modeling approach, and then each analysis in turn.

Design and Analytic Samples

Our panel comprises three nationally representative waves fielded across two presidential cycles.[2] Each wave asked respondents about their confidence in a wide range of institutions, along with their support for democracy and a battery of democratic norms.

Our analyses draw on two samples. The measurement-comparability and party-gap analyses (Analysis III and Analysis IV) are restricted to respondents identifying with the two major parties, since their inferential tasks are (i) establishing whether the confidence battery supports a comparable latent scale within a wave and (ii) estimating party differences on that shared scale. Across waves, the Democrat/Republican sample sizes are Wave 1 ($N_{\text{Dem}} = 2286$, $N_{\text{Rep}} = 1521$), Wave 2 ($N_{\text{Dem}} = 1810$, $N_{\text{Rep}} = 1152$), and Wave 3 ($N_{\text{Dem}} = 1701$, $N_{\text{Rep}} = 1187$). The factor-structure and over-time analyses (Analyses I

[2]Wave 1 in June/July 2018, Wave 2 in August/September 2021, and Wave 3 in October 2023.

and II, respectively) and the downstream analysis (Analysis V) all use the full panel, including Independents; Analysis V additionally estimates effects separately within Democrat, Independent, and Republican strata. Each wave carries cross-sectional survey weights, but our main models are estimated unweighted. We re-estimate the measurement models with our survey weights and report the comparison in Appendix J.

Panel retention is substantial but incomplete: 2,293 respondents completed Waves 1 and 2, 1,888 completed Waves 2 and 3, 1,500 completed Waves 1 and 3, and 1,240 completed all three. The downstream analysis (Analysis V) is estimated on the overlapping subsample for each lag window, with the primary Wave 1→Wave 3 window drawing on the 1,500 respondents observed in both. Attrition is not random, as respondents retained through all three waves are, at baseline, markedly older than those who drop out (a mean age near 55 versus 44) and modestly more educated. To gauge how much this selection matters, we re-estimate the downstream (Analysis V) models using inverse-probability-of-retention weights and find the estimates are substantively unchanged (see Table 10 in Appendix J).

Measures: Confidence in Public and Private Authority

Each wave asks respondents how much confidence they have in each of nineteen institutions, recorded on a four-category ordered scale: "A great deal of confidence," "Only some confidence," "Hardly any confidence," or "No confidence." Consistent with our conception of institutional confidence as legitimacy across public and private authority, we model these items as two substantive domains. The **STATE** domain (ten items) covers political and governing authority: the executive, Congress, political parties, the courts, state government, local government, the FBI, the military, the national press, and local police. The **PRIVATE** domain (nine items) covers civic and market authority: Facebook, Google, Amazon, major companies, banks, colleges and universities, nonprofits, labor unions, and religious organizations.

We include this domain structure, rather than treating confidence as purely unidimensional, for a statistical and a substantive reason. Statistically, it is a more conservative choice to permit variance to be accorded to the two domains. So specified, a unidimensional,

pure-G model would be a special case where there is no variance explained by the domain structure. Where the domains do sop up variance, keeping them avoids the misspecification that forcing a single dimension would introduce. Substantively, the public/private partition is the distinction our theory is built on, and modeling it explicitly allows us to ask whether the structures change over time. We fix the split on theoretical grounds rather than estimating domains from the data, so that change over time reflects shifting attitudes rather than a moving definition of the domains.

Approach: Ordered-Categorical Bifactor CFA

Our confidence responses are ordered categories rather than interval data, so we fit ordered-categorical confirmatory factor analysis (CFA) models in which each observed response $y_{ij} \in \{1, 2, 3, 4\}$ is generated by an underlying continuous propensity y^*_{ij} cut by thresholds τ:

$$y_{ij} = k \;\text{ if }\; \tau_{k-1} < y^*_{ij} \leq \tau_k, \qquad k = 1, \ldots, 4.$$

We estimate the bifactor structure introduced above as an ordered-categorical CFA (Reise 2012; Koch et al. 2018), fit in lavaan (Rosseel 2012) by Weighted Least Squares with Mean and Variance Adjustment (WLSMV), using a probit link and the θ parameterization. This is the standard approach for ordinal indicators with potentially skewed response distributions (Muthén 1984; Flora and Curran 2004). Every item loads on the general factor G and on exactly one domain factor (STATE or PRIVATE). We hold G orthogonal to both domain factors while letting the two domains correlate, and we standardize the latent factors so that estimated differences can be read in standard-deviation units.

Following standard practice (Brown 2015), we free a small number of residual covariances for item pairs whose responses share content beyond the general and domain factors, so that this local dependence is not misattributed to the latent factors. We do so for the three strongest such pairs:

$$\text{Congress} \leftrightarrow \text{Parties}, \qquad \text{Google} \leftrightarrow \text{Amazon}, \qquad \text{Local police} \leftrightarrow \text{Military}.$$

Analysis I: Establishing a Dominant General Factor

Analysis I asks whether the confidence battery is organized by a dominant general factor or only by sector-specific structure. Within each wave, we compare the bifactor model to a correlated two-factor alternative that loads items onto the STATE and PRIVATE domains alone, with no general factor. This comparison tells us whether a general dimension is needed to account for the data or whether the two domains suffice. We summarize fit with the Comparative Fit Index (CFI), the Root Mean Square Error of Approximation (RMSEA), and the Standardized Root Mean Square Residual (SRMR). Beyond fit, we quantify how much of the common variance the general factor carries relative to the domains, using explained common variance (ECV) from the bifactor solution (Rodriguez et al. 2016). ECV(G) is the share of common variance attributable to G, and ECV(STATE) and ECV(PRIVATE) the shares attributable to each domain. Because we fit the model separately in each wave, we can also track ECV(G) across waves. Its trajectory is a central object of our over-time results.

Analysis II: Assessing Generalized Confidence Over Time

Comparing generalized confidence across waves is only meaningful if the items measure the same thing in the same way each time. We guard against possible measurement change by estimating over-time change on a harmonized common-item core, i.e. the confidence items asked with comparable wording and response options in every wave. In our data, this core comprises the nineteen institutions split into the two domains above.[3]

On this core we fit a single ordered-categorical bifactor CFA across all three waves at once (a multi-group model with wave as the grouping variable), retaining the same general and domain factors and the same theory-motivated residual covariances in each wave. In order for us to be able to interpret differences in the latent mean of G as real change, the items must relate to the latent factors the same way in every wave. We therefore impose invariance in the standard sequence, targeted on the general factor: from a configural baseline

[3]In each wave of the survey, we asked about additional institutions not included in the other waves, but do not use those institutions in this analysis, or this paper more generally. These institutions were mostly media-related, and included Fox News (added in Wave 2), major TV network news (Wave 2), CNN (Wave 3), and MSNBC (Wave 3).

(same structure, parameters free), to metric constraints equating the G loadings across waves (metric-G invariance), to scalar constraints additionally equating the item thresholds (scalar-G invariance) (Meredith 1993; Millsap 2011; Steenkamp and Baumgartner 1998). We leave the domain loadings free across waves, since our interest is the comparability of G rather than of the sector-specific structure. We judge each constraint by the change in approximate fit as it is imposed, treating it as supported when $\Delta\text{CFI} \geq -0.010$ and $\Delta\text{RMSEA} \leq 0.015$ (Cheung and Rensvold 2002; Chen 2007; Rutkowski and Svetina 2014). These scalar constraints hold with no thresholds needing release, so wave-to-wave differences in the mean of G reflect real change rather than shifts in how respondents use the response categories (Vandenberg and Lance 2000).

With scalar comparability established, we anchor the scale by fixing the Wave 1 mean of G to zero, so that the Wave 2 and Wave 3 means are interpretable as deviations from 2018 in Wave 1 standard-deviation units. We estimate the wave-to-wave differences (e.g., Wave 3 minus Wave 1) as functions of the model parameters, with standard errors from the model-implied covariance matrix. The resulting estimate of over-time change rests on the same items measured the same way in every wave; we report the latent mean of G by wave, with the Wave 3–Wave 1 contrast and its uncertainty, in the Results.

Analysis III: Within-Wave Party Comparability and the Partisan Gap

Analysis III asks whether Democrats and Republicans can be placed on the same latent confidence scale within each wave, and if so, how far apart they sit. We apply the invariance framework discused above for Analysis II, now comparing the two parties within each wave rather than waves in the pooled sample: from a configural baseline to metric-G (equal G loadings) to scalar-G (equal thresholds) constraints across parties, with the domain loadings again left free, and judged by the same fit-change criteria used in Analysis II.[4] The scalar-G constraints are supported in each wave, with no thresholds freed, so partisan differences in

[4] As before, "invariance" means the items relate to the latent factors the same way across groups, so that group differences reflect differences in latent confidence rather than in measurement.

G can be read as real differences in generalized confidence rather than differences in how Democrats and Republicans use the response scale.

With comparability established, we estimate the Democrat minus Republican difference in the latent mean of G within each wave, fixing one party's mean to zero so the contrast is read in latent standard-deviation units. This within-wave partisan gap, and its growth across waves, is the focus of our party analysis. The MIMIC screens of Analysis IV discussed below then localize which specific institutions carry partisan separation beyond the general factor.

Analysis IV: Identifying Residual Politicization Using MIMIC DIF Screens

Establishing that Democrats and Republicans can be compared on G does not tell us *which* institutions remain partisan flashpoints once generalized confidence is accounted for. To map that residual politicization, we use a multiple-indicator multiple-cause (MIMIC) model to screen for differential item functioning (DIF) (Jöreskog and Goldberger 1975).

A MIMIC model adds a predictor to the measurement model. The items still measure the latent factors as before, but party identification is now allowed to predict where a respondent falls on those factors. In the baseline specification, party predicts G, which is simply how the overall partisan gap is represented. As an example, a Republican with lower G is expected to report lower confidence in *every* institution, in proportion to that institution's loading on G. Differential item functioning is the departure from that pattern. For each institution, we add a *direct* path from party to the item, over and above the item's dependence on G and its domain factor, so that its latent response propensity becomes

$$y_i^* = \lambda_{G,i}\, G + \lambda_{D,i}\, \text{Domain}_i + \gamma_i\, \text{Party} + \varepsilon_i.$$

Under measurement invariance, $\gamma_i = 0$, and two respondents with the same G, one a Democrat and one a Republican, are expected to answer the item identically. A nonzero γ_i means they are not, and that the institution is rated differently along party lines beyond what generalized confidence explains. The general partisan gap thus travels through G, while γ_i

isolates the surplus specific to institution i; we estimate and test it for every institution in every wave (Woods 2009; Jessee 2021). Because γ_i shifts the item's location rather than its loading, this is a screen for *uniform* DIF, or a systematic partisan offset on the item, rather than for party differences in how strongly the item reflects G.

The screen complements Analysis III. There, we use the invariance tests to establish that party comparisons of G are valid at the level of the scale, while here, we use the MIMIC screens to localize the specific institutions that carry partisan separation beyond it.

Analysis V: Downstream Consequences of Generalized Institutional Confidence

Having spent most of the paper analyzing the existence of a general factor, any possible trends in its strength, as well as sub-group differences, we finally turn to consequences it may have beyond institutional ratings alone. To do that, we estimate cross-lagged panel models, by ordinary least squares, that predict democratic norm commitment from prior-wave G. Each model regresses an outcome measured at a later wave on G at an earlier wave, controlling for the respondent's baseline level of that same outcome, party identification, age, education, and gender:

$$\text{outcome}_{it2} \;=\; \alpha + \beta\, G_{it1} + \gamma\, \text{outcome}_{it1} + \delta' \mathbf{X}_{it1} + \varepsilon_{it}.$$

By including the lagged outcome on the right-hand side, the coefficient β captures whether starting higher on G predicts *where a respondent ends up* on democratic norms relative to where they started. That is, we are capturing a conditional-change estimand over time, not a cross-sectional association. Because each respondent contributes a single observation to a given window's regression, we report conventional standard errors.

We examine three lag windows that exploit the panel's three-wave structure. The primary window is Wave 1→Wave 3 (2018→2023), the longest available horizon; the two supplementary windows are Wave 1→Wave 2 (2018→2021) and Wave 2→Wave 3 (2021→2023). Because the democratic-norm items were re-administered in every wave, each window uses wave-appropriate measurements of both baseline and outcome, capturing genuine change

rather than reused Wave 1 values. The analytic sample for each window is the set of respondents observed in both of its waves with complete data on the model's variables. That means roughly 1,500 for the primary Wave 1→Wave 3 window, and somewhat larger for the shorter windows (about 2,300 for Wave 1→Wave 2 and 1,900 for Wave 2→Wave 3), varying modestly across outcomes with item nonresponse.

Our outcome battery consists of fifteen democratic norm items in four clusters. The *Democratic Attitudes* cluster (five items) captures diffuse orientations toward the system: satisfaction with democracy, trust in the federal government, low corruption cynicism, political efficacy, and confidence in the public's wisdom. The *Partisan Norms* cluster (four items) measures cross-party tolerance: rejection of own-party rule-breaking, opposition-party legitimacy, the sense that the opposition is not a serious threat, and willingness to compromise. The Foa and Mounk (2016) *Deconsolidation* cluster (four items) measures rejection of nondemocratic alternatives: a democratic system, a strong leader, military rule, and rule by experts. A final *Democratic Principles* cluster (two items) captures commitment to the formal rules of the game: support for universal suffrage and the view that the president must obey the law. All outcomes are rescaled to $[0, 1]$ with higher values indicating more pro-democratic orientations; complete item wording, response scales, and coding are given in the appendix (Table 5).

As a robustness check on the cross-lagged design, we also compute Pearson correlations between within-person change in G (ΔG) and within-person change in each outcome (Δ outcome) across each lag window. This change-score approach makes no assumptions about baseline levels and asks simply whether respondents whose G rose also became more committed to democratic norms; unlike the cross-lagged regression, it operates entirely within persons. Finally, we examine whether lagged-G effects vary by party identification by re-estimating the primary regression separately within Democrat, Independent, and Republican strata, dropping party identification from the control set within each stratum.

Results

Analysis I: Establishing a Dominant General Factor

We begin by assessing whether institutional confidence behaves primarily as evaluations of two distinct sectors (STATE versus PRIVATE) or as a broader legitimacy-like judgment that spans both. Table 1 compares a correlated two-factor model to an ordered-categorical bifactor specification by wave. In all three waves, the bifactor model fits substantially better (CFI near 0.99; RMSEA near 0.06), indicating that responses are not well captured by two correlated domains alone. Instead, a single generalized confidence factor accounts for a large share of covariation across the entire roster of institutions, with the domain factors capturing structured deviations.

Table 1: Model Fit by Wave

Wave	Model	CFI	RMSEA	SRMR
1	Two-factor (STATE, PRIVATE)	0.940	0.131	0.104
1	Bifactor (G + domains)	0.989	0.060	0.045
2	Two-factor (STATE, PRIVATE)	0.966	0.122	0.096
2	Bifactor (G + domains)	0.992	0.062	0.045
3	Two-factor (STATE, PRIVATE)	0.968	0.117	0.092
3	Bifactor (G + domains)	0.993	0.058	0.044

Notes: CFI = Comparative Fit Index; RMSEA = Root Mean Square Error of Approximation; SRMR = Standardized Root Mean Square Residual. Higher CFI and lower RMSEA/SRMR indicate better fit. All models are estimated by WLSMV with a probit link.

Fit improvements alone cannot determine whether G is substantively dominant. Table 2 reports explained common variance (ECV), which distributes common variance across G and the domain factors. ECV(G) rises from 0.199 in Wave 1 to 0.470 in Wave 3, implying that nearly half of the shared variation in institutional ratings in the final wave is attributable to a single generalized confidence factor. The domain factors remain meaningful, but their relative share declines over time. This is consistent with an environment in which respondents increasingly apply a general legitimacy-like lens across both public and private authority.

This rise in G-dominance is driven by two factors both working in the same direction.

Table 2: Explained Common Variance (ECV) - Bifactor Model

Wave	ECV(G)	ECV(STATE)	ECV(PRIVATE)
1	0.199	0.450	0.350
2	0.406	0.330	0.264
3	0.470	0.296	0.234

Notes: ECV(G) is the share of common variance attributable to the general factor; ECV(STATE) and ECV(PRIVATE) are the corresponding shares attributable to the domain factors (Rodriguez et al. 2016).

First, institutions bind more tightly to the general factor over the panel. As Figure 1 shows, in 2018 most institutions load only weakly on G, a sign of how little a common dimension yet organized the battery, whereas by 2023 most load strongly and positively.[5] This item-level strengthening is the flip side of the aggregate rise in explained common variance reported above (Table 2). As institutions bind more tightly to G, the share of each one's variance attributable to it climbs steadily across the panel. Second, and in tandem, the sector-specific structure recedes. As these institutions are drawn into G, their domain-specific loadings shrink (Appendix Figure 7), so the STATE and PRIVATE factors explain steadily less of the common variance. That the institutions moving most are disproportionately the politically contested ones is consistent with partisan sorting pulling once-distinct evaluations onto a single dimension.

Analysis II: Generalized Confidence over Time

To assess over-time change, we fit the model on a harmonized common-item core and compare wave means on a comparable scale with Wave 1 anchored at 0. Because the scalar-G constraints established above hold across waves, these means are comparable: differences reflect real movement in generalized confidence rather than shifts in how respondents used the response scale. Figure 2 shows a sharp decline from Wave 1 to Wave 2 that then levels off. Relative to Wave 1, the mean falls to -0.294 SD in Wave 2 (SE = 0.029; 95% CI $[-0.352, -0.236]$) and stands at -0.267 SD in Wave 3 (SE = 0.029; 95% CI $[-0.323, -0.210]$).

[5] Because the general factor is weakly identified in 2018 (ECV = 0.20), the individual Wave 1 loadings, including the negative ones, are imprecisely determined and are best read as evidence that no common dimension yet organized the battery, rather than as precise institution-specific estimates.

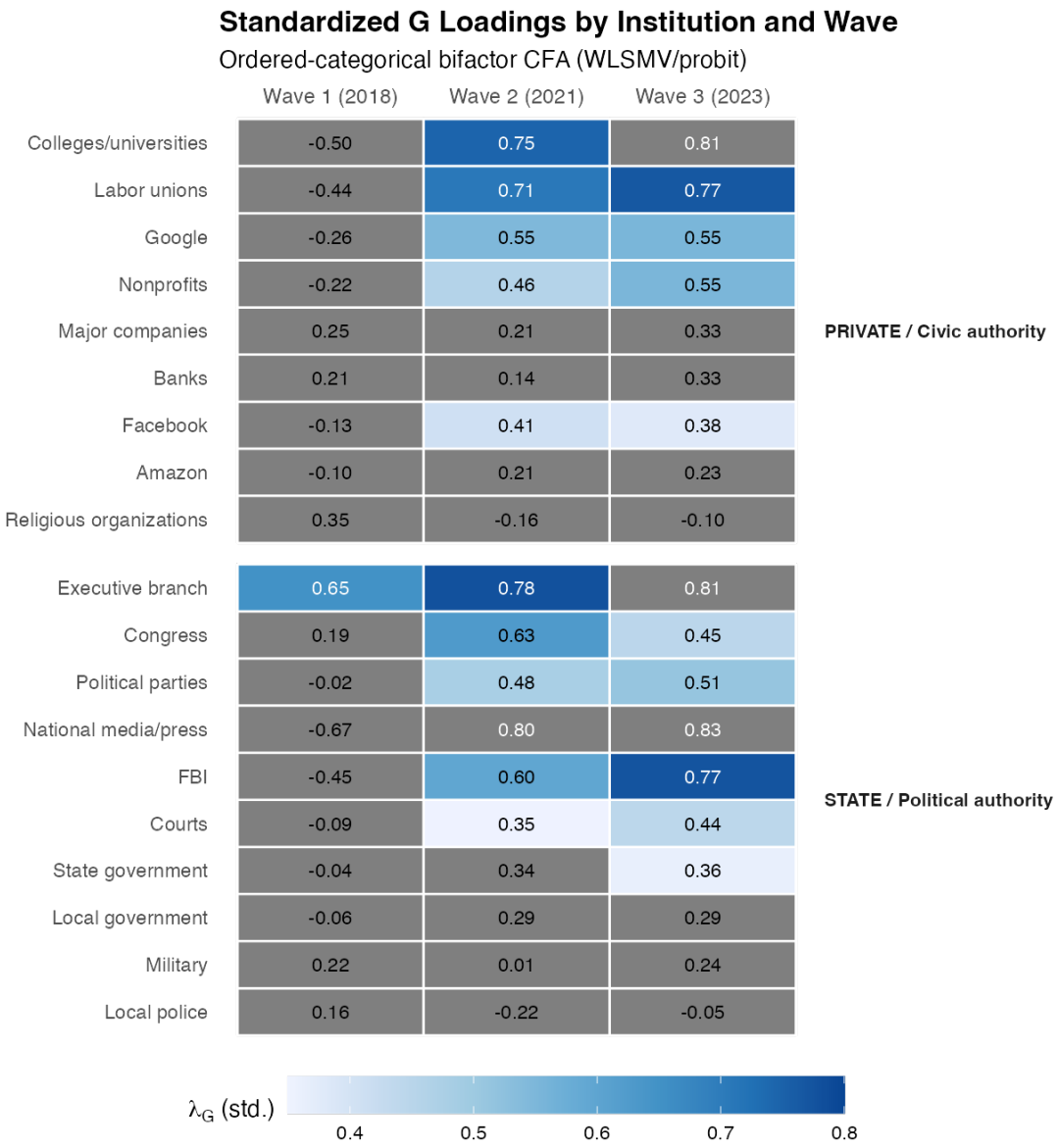

Figure 1: Standardized general-factor (G) loadings by institution and wave, common-item core bifactor model.

These results represent a small-to-moderate downward shift between 2018 and 2023 that was already fully present by 2021. This decline is real but modest, and it is dwarfed by the partisan divergence we turn to next, where the major movement over this period is less a uniform collapse of confidence than a redistribution of it across the parties.

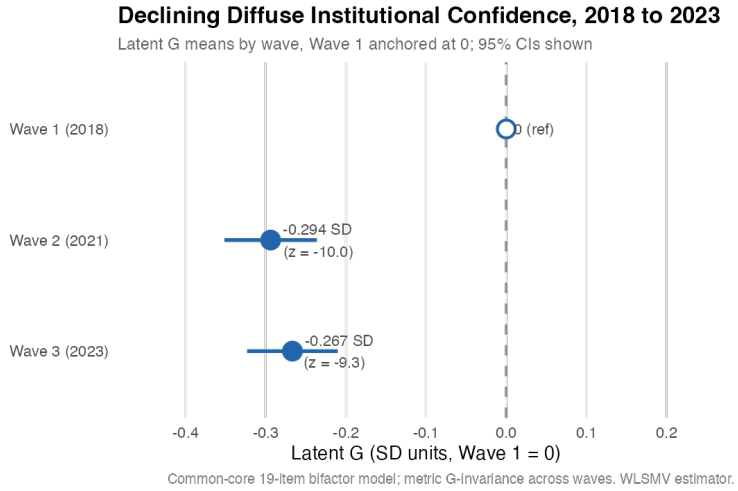

Figure 2: Latent G means by wave on the common-item core, Wave 1 anchored at 0 (95% CIs).

Analysis III: Within-Wave Party Comparability

In order to get there, we must first evaluate whether Democrats and Republicans can be compared on a common scale *within each wave*. Table 3 summarizes the multi-group invariance tests. Configural models fit well, and imposing our metric-G constraints produces only small changes in approximate fit. Scalar-G constraints also satisfy standard criteria in all waves, supporting the interpretation that partisan differences in latent G reflect differences on a comparable response scale rather than threshold drift or differential interpretation (Meredith 1993; Millsap 2011; Chen 2007).

Given scalar comparability, Figure 3 shows latent differences in G, with Republicans as the within-wave reference group. Democrats score higher in every wave, and the gap grows across the period: 0.382 SD in Wave 1 (SE = 0.034; 95% CI $[0.315, 0.448]$), 0.671 SD in Wave 2 (SE = 0.040; 95% CI $[0.592, 0.749]$), and 0.786 SD in Wave 3 (SE = 0.040; 95% CI $[0.707, 0.865]$). Because G spans the full roster of our institutions, these differences are best

Table 3: Within-wave measurement invariance by party (Democrat vs. Republican).

Wave	Config. CFI	Config. RMSEA	ΔCFI (Metric-G)	ΔRMSEA (Metric-G)	ΔCFI (Scalar-G)	ΔRMSEA (Scalar-G)
1	0.990	0.057	-0.004	0.007	0.000	-0.003
2	0.994	0.052	-0.002	0.007	0.000	-0.002
3	0.992	0.056	-0.003	0.006	0.000	-0.003

Notes: Metric-G constrains only G loadings equal across party; domain loadings are free. Scalar-G additionally constrains ordered thresholds equal across party. Traditional guidelines for plausible invariance are $\Delta\text{CFI} \geq -0.010$ and $\Delta\text{RMSEA} \leq 0.015$ (Cheung and Rensvold 2002; Chen 2007; Rutkowski and Svetina 2014).

interpreted as broad-based gaps in generalized legitimacy rather than disagreement about any one institution.

Because thresholds vary across items, latent SD units do not translate mechanically into a fixed number of response categories. The magnitudes here are therefore best interpreted as *broad shifts across many items simultaneously.* In that sense, partisan gaps of 0.38 to 0.79 SD reflect substantial differences in generalized confidence across the institutional roster.

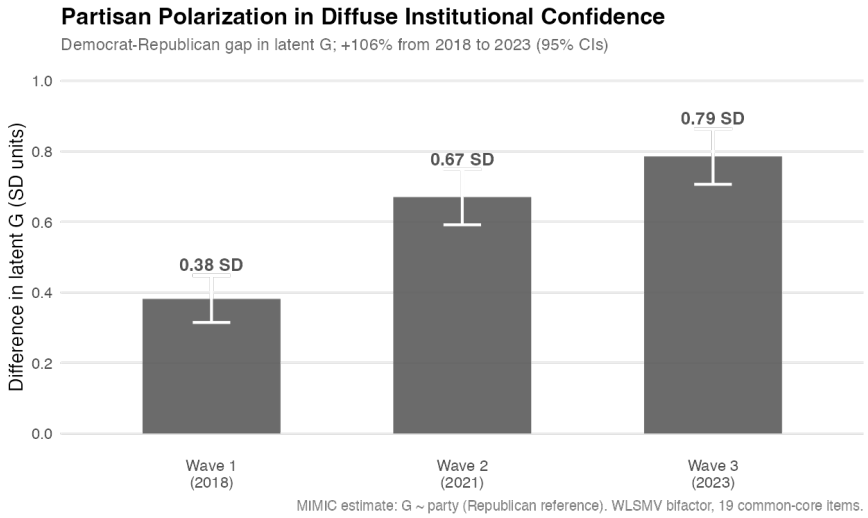

Figure 3: The partisan gap in latent G (Democrat minus Republican) by wave, with 95% CIs.

These latent gaps are visible in raw confidence levels as well. Figure 4 shows the Democrat–Republican gap in mean confidence for each institution in Waves 1 and 3: the gaps are largest, and grow most, for exactly the high-salience institutions that the DIF screens below flag as most politicized.

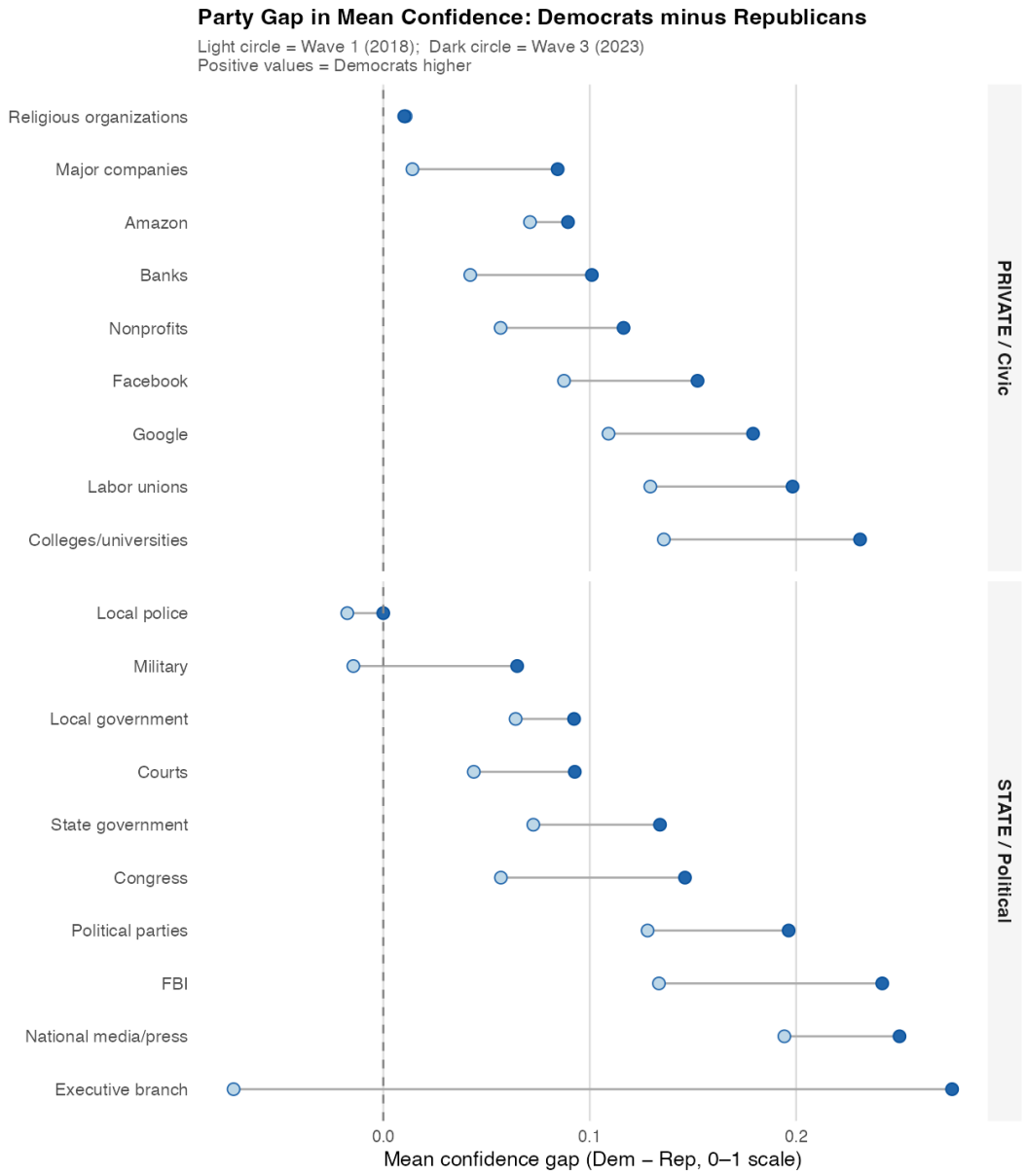

Figure 4: Partisan gap in mean confidence by institution, Wave 1 (2018) versus Wave 3 (2023), sorted by the Wave 3 gap. Raw item means on a 0–1 scale.

Analysis IV: Residual Politicization

We use MIMIC DIF screens to identify institutions whose partisan differences persist beyond the generalized confidence gap and domain structure. Figure 5 plots the direct effect of party on a curated set of items, conditional on G and the domain factors, across the three waves (full item-level results in Appendix Table 7). Two patterns stand out. First, residual partisan sensitivity concentrates in a consistent set of high-salience institutions like the national media, colleges and universities, political parties, and the FBI, each carrying a large Democrat-favored offset in every wave, over and above the general gap in G. These institutions remain distinctive sites of partisan contestation regardless of the political moment.

Second, a subset of institutions reverses partisan valence with control of government. The executive branch draws a Republican-favored residual in 2018 ($\hat{\beta} = -0.234$), under a Republican administration, but large Democrat-favored ones by 2021 and 2023 ($\hat{\beta} = 0.813$ and 0.929), after the change in administration; confidence in the military shifts in the same direction ($-0.065 \rightarrow 0.241$), while local police remain Republican-favored throughout. Residual DIF, then, is not a fixed partisan lean attached to each institution but partly tracks which coalition currently controls or identifies with it. Polarization in institutional confidence thus has both a broad component (i.e. the generalized gap in G), and a concentrated, seemingly incumbency-driven component in high-salience institutions.

Analysis V: G as a Predictor of Democratic Norm Commitment

In Analysis V, we test whether respondents who begin with higher generalized confidence become more committed to democratic norms in later years. We estimate cross-lagged OLS models with each later-wave outcomes regressed on earlier-wave G, the baseline outcome, party identification, and demographics across three lag windows, and corroborate them with within-person change-score correlations. We group the fifteen outcomes into four families: democratic attitudes, partisan norms, resistance to authoritarian alternatives, and core democratic principles. Figure 6 summarizes the primary Wave 1→Wave 3 estimates.

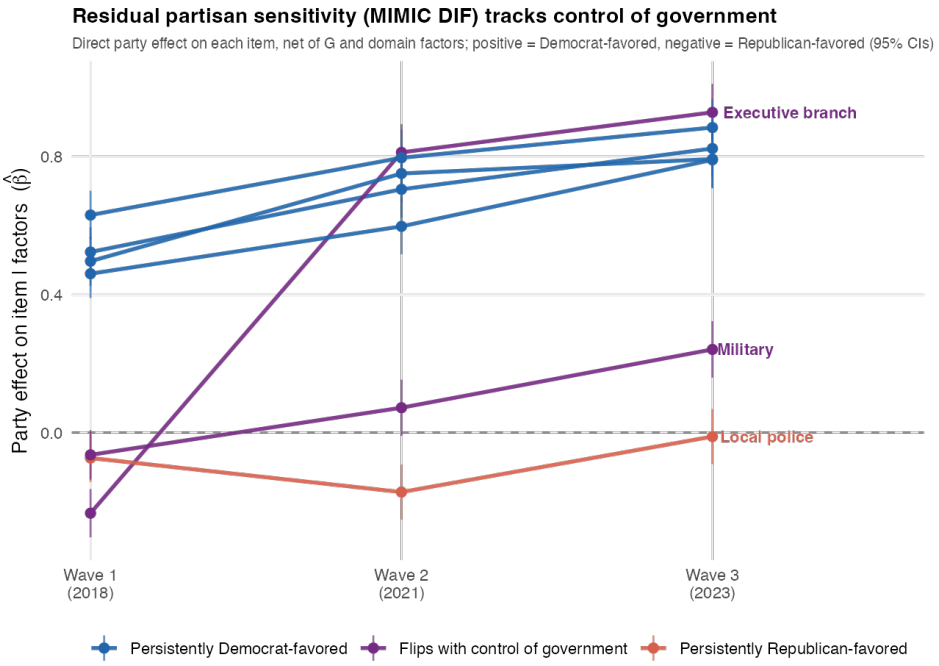

Figure 5: Residual partisan sensitivity (MIMIC DIF) for selected institutions across waves. Each line is the direct effect of party on the item conditional on G and the domain factors; positive values indicate higher confidence among Democrats, negative among Republicans (95% CIs). The stable set in blue (national media, colleges and universities, political parties, and the FBI) is Democrat-favored throughout, while the executive branch and the military reverse valence with the change in administration and local police remain Republican-favored.

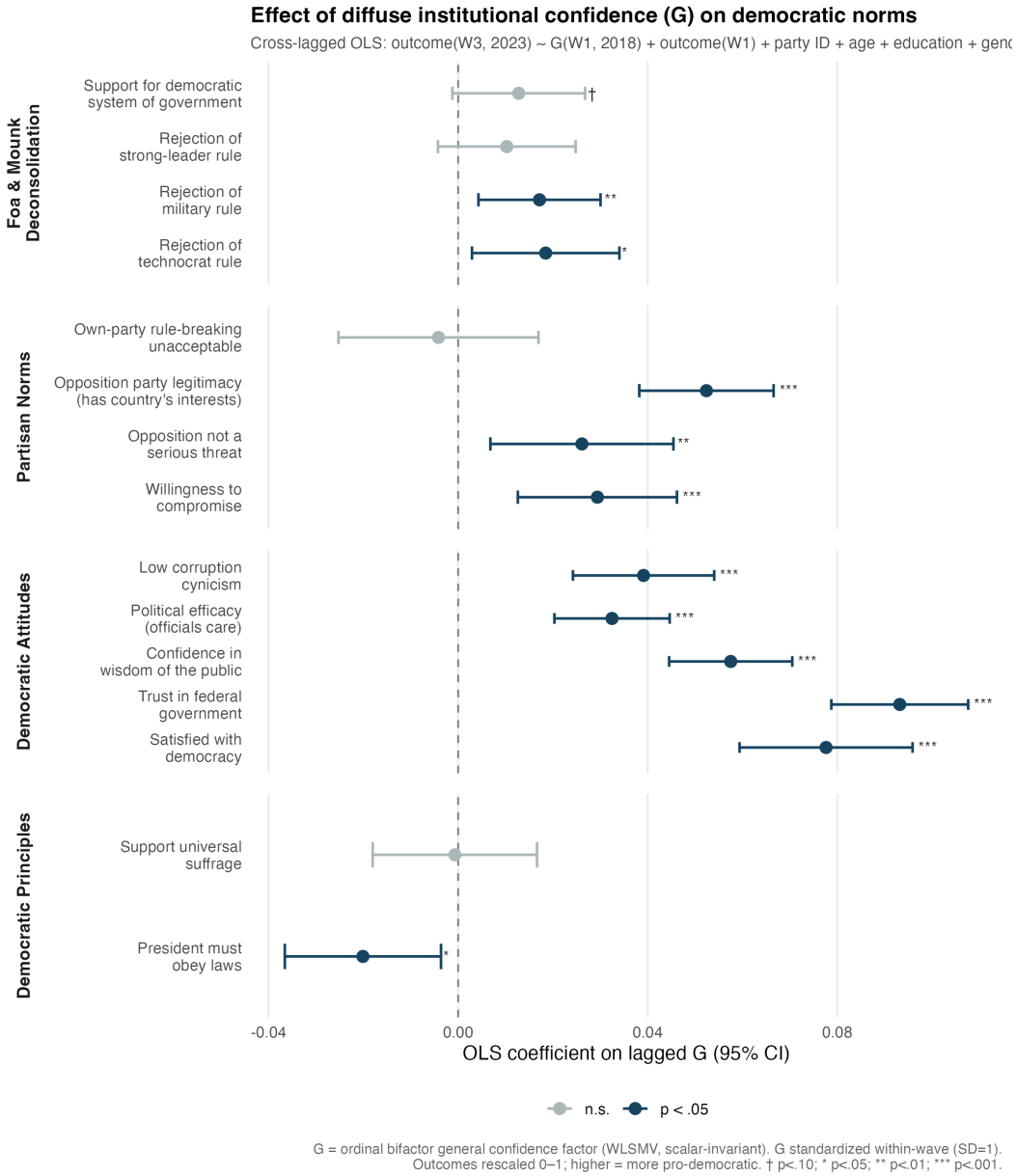

Figure 6: Effect of lagged generalized institutional confidence (G) on democratic norm outcomes (Wave 1→Wave 3, 2018–2023). Points show OLS coefficients; horizontal bars show 95% confidence intervals. Colored points indicate $p < .05$; gray points are non-significant. All outcomes rescaled 0–1 (higher = more pro-democratic). Controls: baseline outcome level, party identification, age, education, and gender. G standardized within-wave (mean 0, SD 1). † $p < .10$; ∗ $p < .05$; ∗∗ $p < .01$; ∗∗∗ $p < .001$.

Democratic attitudes: the Eastonian outcomes

The five Democratic Attitudes outcomes (trust in the federal government, satisfaction with democracy, efficacy, corruption cynicism, and confidence in the wisdom of the public) are significantly predicted by G in all three lag windows and both analytical designs. This cluster represents the outcomes most directly aligned with Eastonian diffuse support: respondents who enter the study period with higher generalized institutional confidence end up, five years later, with meaningfully higher commitment to democratic governance, net of where they began.

The estimates are substantively large. In the primary Wave 1→Wave 3 window, a one-standard-deviation increase in G at Wave 1 (2018) predicts a 9.3 percentage-point increase in government trust at Wave 3 (2023; $b = .093$, SE $= .007$, $p < .001$, $N = 1{,}500$) and a 7.8 percentage-point increase in democratic satisfaction ($b = .078$, SE $= .009$, $p < .001$, $N = 1{,}500$), with parallel effects for efficacy ($b = .032$), corruption cynicism ($b = .039$), and public wisdom ($b = .058$), all estimated at statistically significant levels. The pattern is stable across all three windows (see Appendix Figure 9).

The change-score analysis confirms the cross-lagged results through an entirely independent design. Across the full Wave 1→Wave 3 panel, the correlation between ΔG and Δ(government trust) is $r = .259$ (95% CI $[.211, .306]$, $N = 1{,}492$). This represents a moderate effect for a five-year panel with substantial exogenous noise from political events, life events, and attrition. The analogous correlations for democratic satisfaction ($r = .170$), efficacy ($r = .165$), corruption cynicism ($r = .192$), and public wisdom ($r = .137$) are consistent and precisely estimated. The convergence of the cross-lagged OLS and change-score results across a broadly similar rank ordering of outcomes shows the signal is robust to the choice of analytic design and strengthens a potential causal claim.

G as a potential proxy for party identification

In Analysis III above, we show that G is strongly stratified by party. This may raise a natural concern that the downstream associations merely reflect the fact that Democrats trust both

institutions and government simultaneously. We use party-stratified regressions to directly test this concern. Among Democrats, Independents, and Republicans separately, the G effect on government trust is $b = .093$ ($p < .001$, $N = 644$), $b = .097$ ($p < .001$, $N = 296$), and $b = .085$ ($p < .001$, $N = 460$), respectively. The estimates are nearly identical across strata. Parallel results hold for democratic satisfaction (Democrat $b = .080$; Independent $b = .077$; Republican $b = .068$, all $p < .05$). A Republican in 2018 who scored one standard deviation higher on generalized institutional confidence ended up, by 2023, meaningfully more committed to government trust and democratic satisfaction than a Republican who scored lower, controlling for where each started. Importantly, G is *not* proxying for party.

Partisan norms as a buffer against affective polarization

Three of four partisan norm outcomes are significantly predicted by lagged G. The most consistent is opposition party legitimacy, that is, the belief that the other party has the country's interests at heart. The effect tied to this norm is significant in all three windows ($b = .036$ W1→W2; $b = .035$ W2→W3; $b = .052$ W1→W3, all $p < .001$) and symmetric across party strata (Democrat $b = .047$; Independent $b = .046$; Republican $b = .050$, all $p < .05$). High generalized confidence predicts viewing the opposing party as more legitimate regardless of which party one belongs to. Effects on threat perception (viewing the opposition as less of a serious threat) and willingness to compromise are null or marginal in the Wave 1→Wave 2 window but strengthen to significance by Wave 1→Wave 3, consistent with deepening polarization over the panel period creating more room for G to serve as a buffer. Own-party accountability (holding one's own party to rule-following norms) is null in all windows, implying that G operates through cross-party tolerance rather than intra-group discipline.

Deconsolidation outcomes and period-specificity

The Foa and Mounk (2016) deconsolidation battery exhibits significant effects in at least one window for three of four items, but the pattern is period-specific. Rejection of military rule and rejection of strong-leader rule are significant in the Wave 1→Wave 2 window

(2018→2021), consistent with the Trump-era salience of these anti-democratic governance alternatives. Rejection of technocrat rule and support for the democratic system show stronger effects in the Wave 2→Wave 3 window (2021→2023), consistent with COVID-era debates about technocratic governance and post-January 6 democratic legitimacy concerns. The deconsolidation effects are concentrated among Democrats and Independents in party-stratified models, presumably reflecting the asymmetric salience of authoritarian governance alternatives during the Trump administration.

No clear effects for core democratic principles

The two core-principle items show no substantive G influence. Support for universal suffrage is essentially null across specifications, with cross-lagged $b = -.001$, $p = .94$ and the only detectable signal a marginal negative change-score correlation over the full panel, $r = -.052$, $p = .045$. Belief that the president must obey laws and courts is likewise near zero, with a small negative cross-lagged coefficient ($b = -.020$, $p = .017$) that we suggest treating cautiously given the item's heavy ceiling (roughly 86% of all respondents agree over the three waves). Substantively, generalized confidence predicts the evaluative and relational components of democratic commitment (trust, satisfaction, tolerance of the opposition), but not baseline procedural principles, which command broad agreement regardless of how much confidence one places in institutions.

Discussion and Conclusions

In this paper, we find that, across five years and nineteen institutions spanning public and private authority, Americans' confidence increasingly collapsed onto a single dimension. Moreover, that dimension became more dominant, more partisan, and more consequential for democratic commitment. In this final section, we briefly review our findings before returning to the theoretical stakes we laid out at the start, considering what the findings imply for scholarship and for politics, and closing with the limits of the evidence and the questions the paper opens. We start with the findings.

We identify five major results. First, institutional confidence is increasingly *one* thing. A bifactor model that lets a general factor and sector-specific factors compete fits far better than sector structure alone, and the share of common variance the general factor G carries rises from roughly a fifth in 2018 to nearly half by 2023. This consolidation runs through two channels working together: institutions bind more tightly to G over time, while the sector-specific structure that once separated public from private evaluations recedes. The institutions doing the most moving are the politically contested ones.

Second, this newly dominant dimension drifted modestly downward in level over time. By the second wave in 2021, it had decreased by nearly three-tenths of a standard deviation, and remained there after the third wave in 2023. Third, the dimension pulled sharply apart by party: the Democrat to Republican gap in G roughly doubled over the panel, from about 0.38 to 0.79 standard deviations. We consider this to be a form of redistribution of confidence across the two parties, instead of a uniform collapse across the board. Fourth, we identify a residual layer of item-specific politicization. Much of it is stable, with the national media, universities, political parties, and the FBI carrying a persistent Democrat-favored tilt and local police remaining Republican-favored throughout, but part of it tracks control of government. When the White House flipped in 2021, so too did the confidence in the executive branch and the military, from Republican in the first wave to Democratic-leaning in the second and third.

Finally, from a downstream perspective, G is consequential. Respondents who began the panel with higher generalized confidence became, five years later, more trusting of government, more satisfied with democracy, and more willing to grant the opposing party legitimacy. These effects are, importantly, *not* partisanship by another name, as they hold at nearly identical magnitudes inside each party. There are limits, however, to what this generalized confidence touches. It predicts some components of democratic commitment but not baseline procedural principles, which command broad agreement regardless of how much confidence one places in institutions.

These findings speak directly to the three questions we posed at the outset. The first set three potential mechanisms for the emergence of a generalized confidence against each

other.[6] Our findings that G became more dominant and more partisan over time support the case for a partisan-sorting story, with parties moving apart rather than together. The incumbency-driven reversals in the executive and the military strengthen the point. Respondent evaluations are tracking whose team controls the institution, not its long-term performance. The dispositional and correlated-shock accounts could still contribute, but neither can explain the asymmetry the way that the partisan sorting mechanism does.

The second question was Eastonian. The increasingly dominant general factor we identify is suggestive of something like Easton's diffuse support. As the common component grows, confidence in any given institution reflects less that institution's recent record and more a standing orientation toward the institutional order as a whole. This adjudicates, at least for the United States over this period, the crisis-versus-fluctuation debate. The common reservoir is real and becoming more influential, which is the signature of a shift in diffuse legitimacy rather than a churn of institution-specific complaints (Valgarðsson et al. 2025; van der Meer 2017). Yet our results put a revealing twist on Easton. Diffuse support was originally theorized as a system-*supporting* reservoir, a store of legitimacy that lets a political order weather disappointment (Easton 1975). What we observe, on the other hand, is a diffuse orientation that has become an object of partisan division. The "reservoir" is split, so that each party increasingly extends or withholds legitimacy to the whole institutional order as a bloc. A generalized legitimacy that is itself polarized is a different, and far more brittle, foundation than the one Easton described.

Our third question was whether any of this matters for democracy, and here we specifically referenced the deconsolidation and democratic-vulnerability literatures. Recent work argues that intense partisanship pushes citizens to resolve conflicts between democratic principle and team loyalty in favor of party (Graham and Svolik 2020; Svolik 2019). Our downstream results suggest generalized confidence is one thing that pulls the other way: within each party, higher G predicts firmer commitment to the norms that restrain partisan conflict.

[6]The three included a general factor that stemmed from a general and consistent disposition of either high or low trust, a general factor that occurred when a host of institutions all simultaneously experienced shocks or failures, and a final factor where partisan sorting becomes strong enough that most or all institutions are viewed through the same partisan lens.

Generalized confidence looks like a genuine democratic resource in the sense Foa and Mounk's agenda invites us to look for (Foa and Mounk 2016), but does not extend to all of our democratic norms.

Outside of our three main questions, it is also important to note that the paper extends the confidence tradition past the boundary of the state. Given we find a single factor that spans the traditional governmental bodies, as well as Amazon and colleges and banks, we provide support and some vindication for the long-standing but occasionally under-theorized practice of asking Americans about private and public institutions in one battery (Lipset and Schneider 1983; Smith 2008). Generalized legitimacy seems to extend well past our traditional limits in the current mixed ecology of authority. In all, these findings and answers lead us to several implications for both academia and the real-world.

For scholarship, the most immediate implication deals with both the conceptualizing of confidence and its measurement. A battery of "trust in institutions" items is not a bag of separate attitudes, nor a proxy for trust in government, but increasingly a single measure of generalized legitimacy that spans multiple sectors. Researchers need to address it in that context, by estimating the general factor directly, within a measurement framework that establishes comparability before interpreting gaps. This not only recognizes the truth of the current dynamic, but also allows confidence researchers to speak to affective polarization, private-governance, and democratic-backsliding literatures that have largely developed apart. Methodologically, the paper offers a template for modeling institutional confidence in polarized settings. We show the importance and benefits of establishing within-group and over-time comparability, separating the general factor from domain structure, and disclosing residual item-level politicization through DIF screens.

The real-world implications cover a wide range of important and current issues in democratic governance. First, our findings suggest that any program to "restore trust" is going to face real obstacles. These campaigns generally focus on particular sectors, with maybe as wide a goal as to increase confidence in government, but this misses the degree to which the dimension includes institutions in other sectors *and* the degree to which it may be party-stratified. As such, these appeals are likely to be read as credible by one side of the aisle

and mere partisan messaging by the other. In short, there is likely less space for broad confidence-building.

Outside of the governing institutions, our findings support what some leaders in arguably non-governmental institutions have already discovered the hard way. Specifically, private and semi-private actors like universities, major companies, and social media platforms now exist inside the broader partisan legitimacy contest rather than outside. Debates over things like the free speech of students, content moderation, or AI data centers are all areas that have become secondary fronts in the same polarized struggle over legitimacy that governmental institutions have dealt with over the past few decades. There is no longer the expectation that these types of institutions can make legitimate policy "internal" to the institution while escaping partisan bickering or interference.

Finally, while we find that generalized confidence can support democratic norms across party lines, this is undermined by the fact that this confidence is unevenly held. Worse yet, finding that confidence in the executive and the military now swings with control of government suggests it is weakest in the coalition that is out-of-power. Generalized confidence is part of what lets citizens accept unwelcome outcomes as legitimate. Polarization within it means that democratic restraint runs thinnest in the coalition being asked to concede, a recipe for a rejection of traditional norms and guardrails, including most importantly the peaceful transfer of power.

We should, of course, be qualified in our conclusions. Our analyses rest on a three-wave panel with partial re-interviews. Retained respondents are older and more educated at baseline than the full Wave 1 sample, cross-sectional weighting does not fully absorb compositional change, and the balanced Wave 1→Wave 3 panel ($N = 1{,}240$) is a selectively retained subsample. Finally, the downstream results in Analysis V, though built on a cross-lagged design that controls for baseline norms and demographics, remain observational. We cannot, for instance, rule out unmeasured characteristics that move both G and democratic-norm trajectories. Still, the findings extending from these analyses are consistent across multiple approaches, and survive a variety of rigorous checks. We believe they warrant additional investigation in a variety of areas.

First, this paper was partially inspired by a debate in comparative politics about the degree to which western democracies are experiencing a general decline in institutional trust versus a series of coordinated shocks across institutions during a time of political upheaval. While that debate is far from settled, it would also be interesting to carry out the types of analyses we complete for this paper in individual countries across the world. One should expect that generalized confidence operates differently across national borders, and the American case is, as always, a very specific case. It may be that the dynamic we identify here is less (or more) strong in other democracies.

Second, and related, there are likely to be contexts in which American generalized confidence varies from the national baseline. These contexts may be geographical (some states or regions may see more or less institutional coherence), demographic (age and gender in particular seem likely to shift the way legitimacy is internalized), or socio-economic in nature. Additional work in these sub-samples of the population is likely to yield a finer view of the structure of generalized confidence in this country.

One final area of great interest is the "mechanism of the mechanism" we identify in this paper. We show that partisan sorting is doing much of the work in strengthening the dominance of a generalized confidence, but little about how that sorting worked in this specific case. Exploration of the media diets of individuals, the elite cues they were exposed to, and the timing of specific institutional controversies are all likely to better inform us about how the sorting was able to be so effective and wide-ranging.

All in all, there remains an open question about whether a unified dimension of confidence will prove to be a foundational aspect of American democracy going forward, and the directions it is likely to push the national project.

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Appendix

Appendix A: Institutional-Confidence Battery

This appendix documents the nineteen-item institutional-confidence battery that forms the common core of every model in the paper: the verbatim item wording, the four-category ordered response scale, and the fixed assignment of each institution to the STATE or PRIVATE domain. It supports the measurement choices described in the Measures subsection of Data and Methods.

Respondents were asked how much confidence they have in each institution, on a four-category ordered scale: "A great deal of confidence," "Only some confidence," "Hardly any confidence," "No confidence." Items are scored so that higher values indicate greater confidence.

Table 4: Institutional Confidence Battery

Institution (verbatim)	Variable	Raw item
STATE / Political authority		
The executive branch of the U.S. government	`ConfExec`	A5
Congress	`ConfCongress`	A8
Political parties	`ConfPolParties`	A11
The courts	`Confcourts`	A7
State government in your state	`ConfStategov`	A12
Local government in your city, town, or county	`ConfLocalgov`	A13
The Federal Bureau of Investigation (FBI)	`ConfFBI`	A10
The military	`Confmilitary`	A9
Local police in your city, town, or county	`ConfLocalPol`	A17
The press	`Confpress`	A6
PRIVATE / Civic authority		
Facebook	`ConfFacebook`	A18
Google	`ConfGoogle`	A19
Amazon	`Confamazon`	A20
Major companies	`ConfMajorComps`	A1
Banks and financial institutions	`ConfBanks`	A2
Colleges and universities	`ConfColleges`	A14
Nonprofits	`ConfNonprofits`	A15
Organized labor	`ConfOrglabor`	A3
Organized religion	`ConfOrgreligion`	A4

Later waves additionally asked about several institutions not fielded in every wave. These

were mostly outlet-specific media (e.g., Fox News, CNN, MSNBC, major TV network news), plus Twitter, the CDC, and environmental organizations, and, in 2018 only, philanthropy and private foundations. Because these were not asked in all three waves, they fall outside the common core and are not used.

Appendix B: Democratic-Norm Outcome Battery

Table 5: Democratic-Norm Outcome Battery

Var.	Item (verbatim)	Response scale	Recode
Foa & Mounk deconsolidation (4)			
`r1`	Having a democratic political system [as a way of governing the country]	1 = Very good … 5 = Very bad way of governing	reversed
`r2`	Having a strong leader	1 = Very good … 5 = Very bad way of governing	kept
`r3`	Having the military rule	1 = Very good … 5 = Very bad way of governing	kept
`r4`	Having experts make decisions	1 = Very good … 5 = Very bad way of governing	kept
Partisan norms (4)			
`w`	It is acceptable for your party to break a few rules to beat the opposing party	1 = Completely acceptable … 4 = Not at all acceptable	kept
`x`	The opposing party has the best interests of the country in mind	1 = All of the time … 4 = Never	reversed
`y`	The opposing party is a threat to the United States	1 = Very serious threat … 4 = Not a threat at all	kept
`z`	Your party should compromise with vs. stick to its positions	1 = Always compromise … 4 = Always stick to positions	reversed
Democratic attitudes (5)			
`ww`	Satisfied with the way democracy is working	1 = Very satisfied … 5 = Very dissatisfied	reversed
`ll`	Trust the government in Washington to do what is right	1 = Just about always … 4 = Almost never	reversed
`i`	How many U.S. politicians are corrupt	1 = Almost none … 5 = Almost all	reversed
`j`	Public officials don't care much what people like me think	1 = Strongly agree … 5 = Strongly disagree	kept
`k`	Confidence in the wisdom of the American people making political decisions	1 = A very great deal … 4 = None at all	reversed
Democratic principles (2)			
`m`	American democracy would be better/worse if only people informed about politics could vote	1 = Much better … 5 = Much worse	kept
`o`	President must obey the laws and courts vs. is not bound by them	1 = Not bound; 2 = Must always obey	0/1

Notes: All items are rescaled to $[0, 1]$ with higher values indicating a more pro-democratic orientation. Don't-know/refused codes $(-1, 8, 9)$ are set to missing. Items marked "reversed" are recoded so higher values are more pro-democratic; `o` is recoded to 0/1 (1 = the president must obey).

Appendix C: Standardized G and Domain Loadings by Wave

This appendix reports the full standardized loadings underlying Analysis I: for each institution and wave, its loading on the general factor G and on its STATE or PRIVATE domain factor. These are the item-level values summarized graphically in the main text (Figure 1). They show that most institutions load weakly, or in some cases negatively, on G in 2018 and strongly and positively by 2023, while the domain loadings decline in magnitude over the same period.

Table 6: Standardized G and domain loadings by institution and wave (common-item core bifactor model).

	General factor (G)			Domain factor		
Institution	W1	W2	W3	W1	W2	W3
STATE / Political authority						
Executive branch	0.65	0.78	0.81	0.57	0.45	0.36
Congress	0.19	0.63	0.45	0.71	0.51	0.56
Political parties	-0.02	0.48	0.51	0.69	0.54	0.53
Courts	-0.09	0.35	0.44	0.76	0.68	0.64
State government	-0.04	0.34	0.36	0.77	0.63	0.64
Local government	-0.06	0.29	0.29	0.77	0.71	0.73
FBI	-0.45	0.60	0.77	0.67	0.58	0.39
Military	0.22	0.01	0.24	0.56	0.71	0.60
Local police	0.16	-0.22	-0.05	0.65	0.77	0.77
National media/press	-0.67	0.80	0.83	0.53	0.31	0.27
PRIVATE / Civic authority						
Facebook	-0.13	0.41	0.38	0.60	0.52	0.53
Google	-0.26	0.55	0.55	0.60	0.45	0.43
Amazon	-0.10	0.21	0.23	0.54	0.59	0.59
Major companies	0.25	0.21	0.33	0.77	0.79	0.75
Banks	0.21	0.14	0.33	0.74	0.79	0.72
Colleges/universities	-0.50	0.75	0.81	0.64	0.39	0.29
Nonprofits	-0.22	0.46	0.55	0.58	0.48	0.40
Labor unions	-0.44	0.71	0.77	0.59	0.28	0.17
Religious organizations	0.35	-0.16	-0.10	0.54	0.60	0.64

Notes: Standardized loadings from the ordered-categorical bifactor model on the harmonized common-item core (WLSMV/probit). Each institution loads on the general factor G and on its designated domain factor (STATE or PRIVATE). Because G is weakly identified in 2018 (ECV = 0.20), the Wave 1 loadings are imprecisely determined.

Appendix D: Domain-Factor Loadings Across Waves

This appendix presents the domain-factor loadings (net of G) by institution and wave. As institutions are drawn into G over the panel, their domain-specific loadings shrink, most sharply for the national media, colleges and universities, labor unions, and the FBI; the loadings remain positive throughout rather than reversing sign.

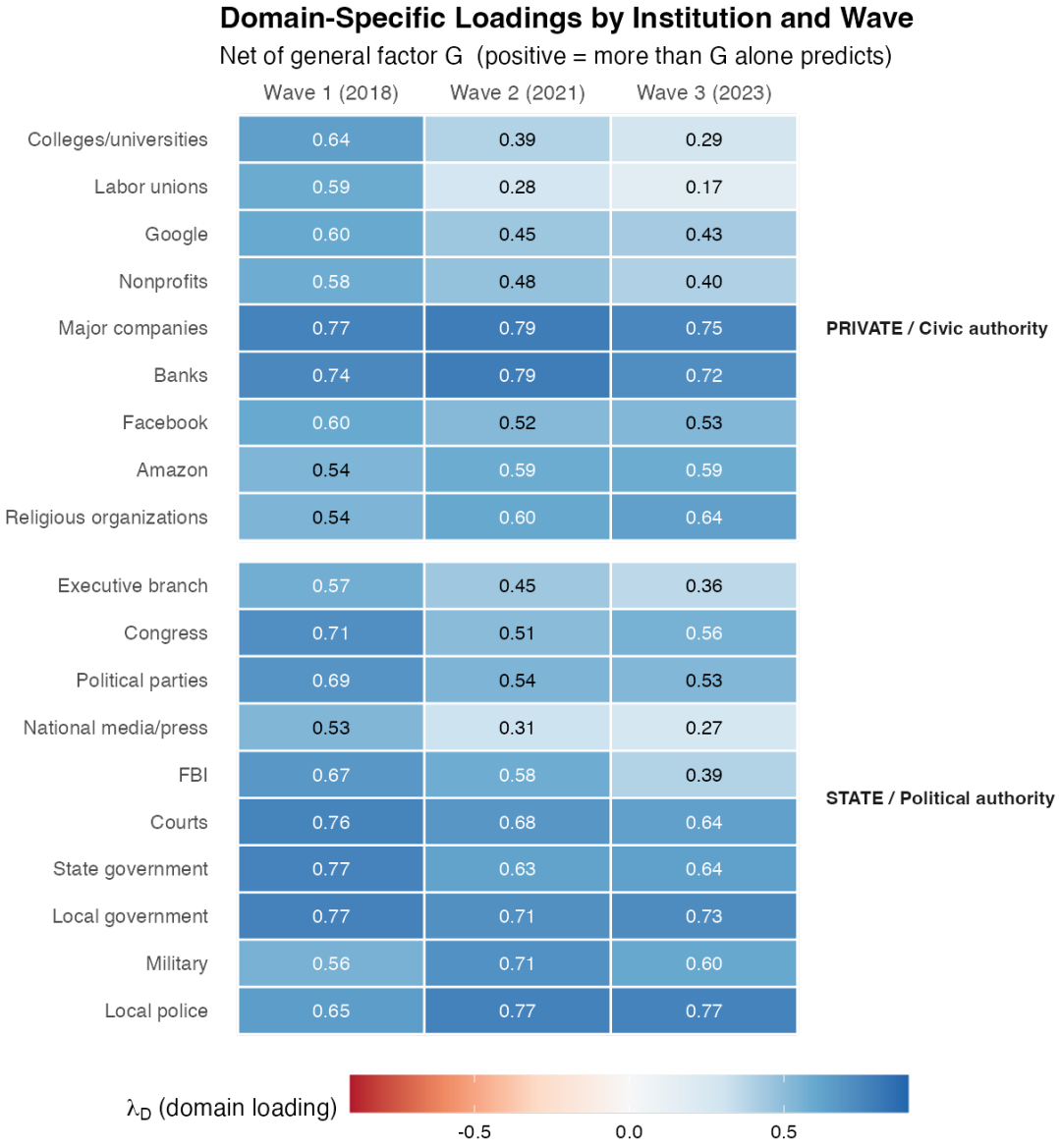

Figure 7: Domain-factor loadings by institution and wave (net of G).

Appendix E: Item-Level Variance Decomposition

This appendix decomposes each item's variance into the share attributable to G, the share attributable to its domain factor, and uniqueness, comparing Wave 1 with Wave 3. It shows at the item level the same shift the ECV trajectory documents in aggregate in Analysis I: the G share rises across the battery, most steeply for the executive, national media, FBI, and colleges.

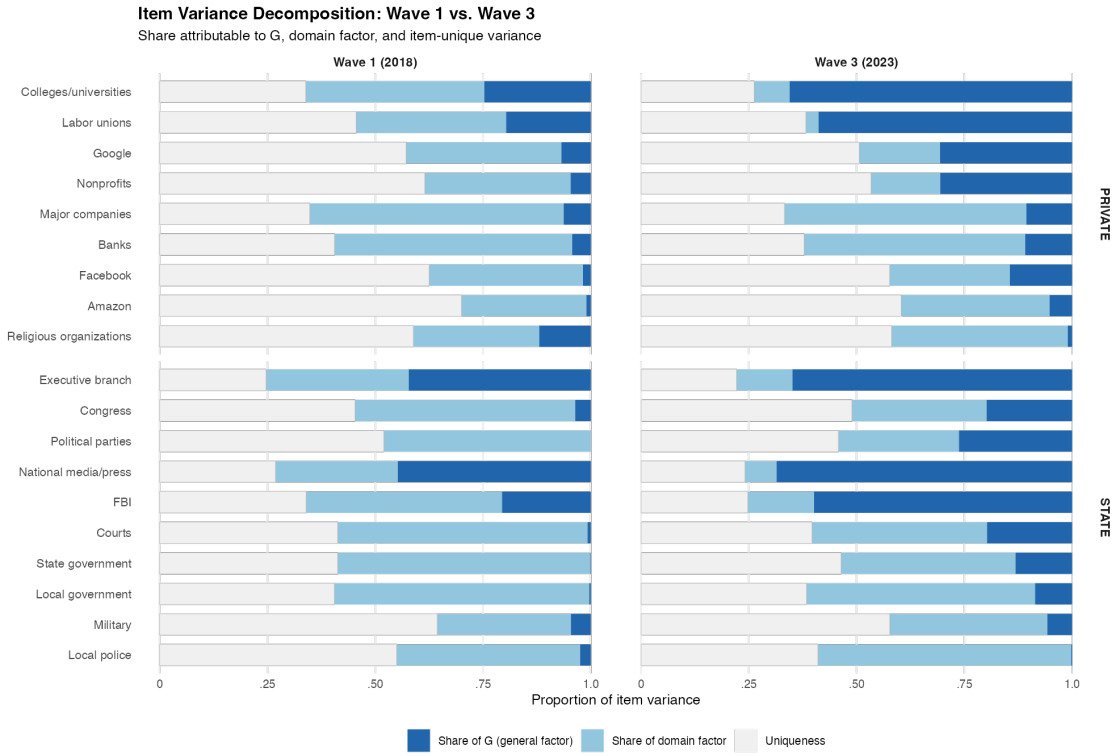

Figure 8: Item-level variance decomposition (share attributable to G, to the domain factor, and uniqueness), Wave 1 versus Wave 3.

Appendix F: Full MIMIC DIF Results

This appendix reports the complete MIMIC differential-item-functioning results: the direct effect of party on every institution, conditional on G and the domain factors, in each wave. It is the full version of the curated set shown in the main text (Figure 5) and documents both the persistently Democrat-favored institutions and the executive and military valence reversals discussed in Analysis IV.

Table 7: Full MIMIC DIF results by wave (items $\sim$ party | factors).

Wave	Item	$\hat{\beta}$	SE	z
1	National media/press	0.631	0.036	17.547
	Colleges/universities	0.524	0.036	14.363
	Political parties	0.497	0.036	13.704
	Labor unions	0.481	0.036	13.337
	FBI	0.461	0.036	12.773
	Google	0.408	0.036	11.230
	Facebook	0.294	0.036	8.165
	Amazon	0.284	0.037	7.754
	State government	0.259	0.036	7.245
	Nonprofits	0.248	0.037	6.733
	Local government	0.245	0.036	6.851
	Executive branch	-0.234	0.036	-6.534
	Congress	0.211	0.036	5.866
	Courts	0.170	0.036	4.716
	Banks	0.162	0.036	4.472
	Local police	-0.074	0.036	-2.075
	Military	-0.065	0.037	-1.761
	Major companies	0.061	0.036	1.687
	Religious organizations	0.033	0.036	0.929

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Wave	Item	$\hat{\beta}$	SE	z
2	Executive branch	0.813	0.042	19.463
	National media/press	0.797	0.042	19.112
	Political parties	0.751	0.042	17.847
	Congress	0.751	0.042	18.009
	Colleges/universities	0.706	0.042	16.879
	Labor unions	0.659	0.042	15.773
	Google	0.610	0.041	14.735
	FBI	0.598	0.041	14.488
	Facebook	0.501	0.042	11.796
	Nonprofits	0.416	0.042	9.903
	State government	0.381	0.041	9.304
	Amazon	0.334	0.041	8.097
	Local government	0.323	0.041	7.829
	Banks	0.297	0.041	7.232
	Courts	0.261	0.041	6.368
	Major companies	0.245	0.041	5.916
	Local police	-0.172	0.041	-4.232
	Military	0.072	0.042	1.732
	Religious organizations	-0.023	0.041	-0.551
3	Executive branch	0.929	0.042	22.060
	National media/press	0.885	0.042	21.121
	Colleges/universities	0.824	0.042	19.396
	Political parties	0.793	0.043	18.642
	FBI	0.791	0.042	18.731
	Labor unions	0.717	0.042	17.114
	Google	0.645	0.042	15.305

Continued on next page

Wave	Item	$\hat{\beta}$	SE	z
	Congress	0.566	0.042	13.511
	Facebook	0.535	0.042	12.668
	Nonprofits	0.492	0.042	11.704
	State government	0.465	0.041	11.306
	Banks	0.385	0.041	9.364
	Local government	0.361	0.042	8.656
	Courts	0.348	0.041	8.437
	Major companies	0.336	0.042	8.100
	Amazon	0.320	0.041	7.808
	Military	0.241	0.042	5.771
	Religious organizations	0.036	0.041	0.886
	Local police	-0.012	0.041	-0.284

Appendix G: Lagged-G Effects Across All Windows

This appendix shows the lagged-G effects on all fifteen democratic-norm outcomes across all three lag windows, extending the main-text figure. It confirms that the Eastonian democratic-attitudes effects are positive and significant in every window, supporting the stability claim in Analysis V.

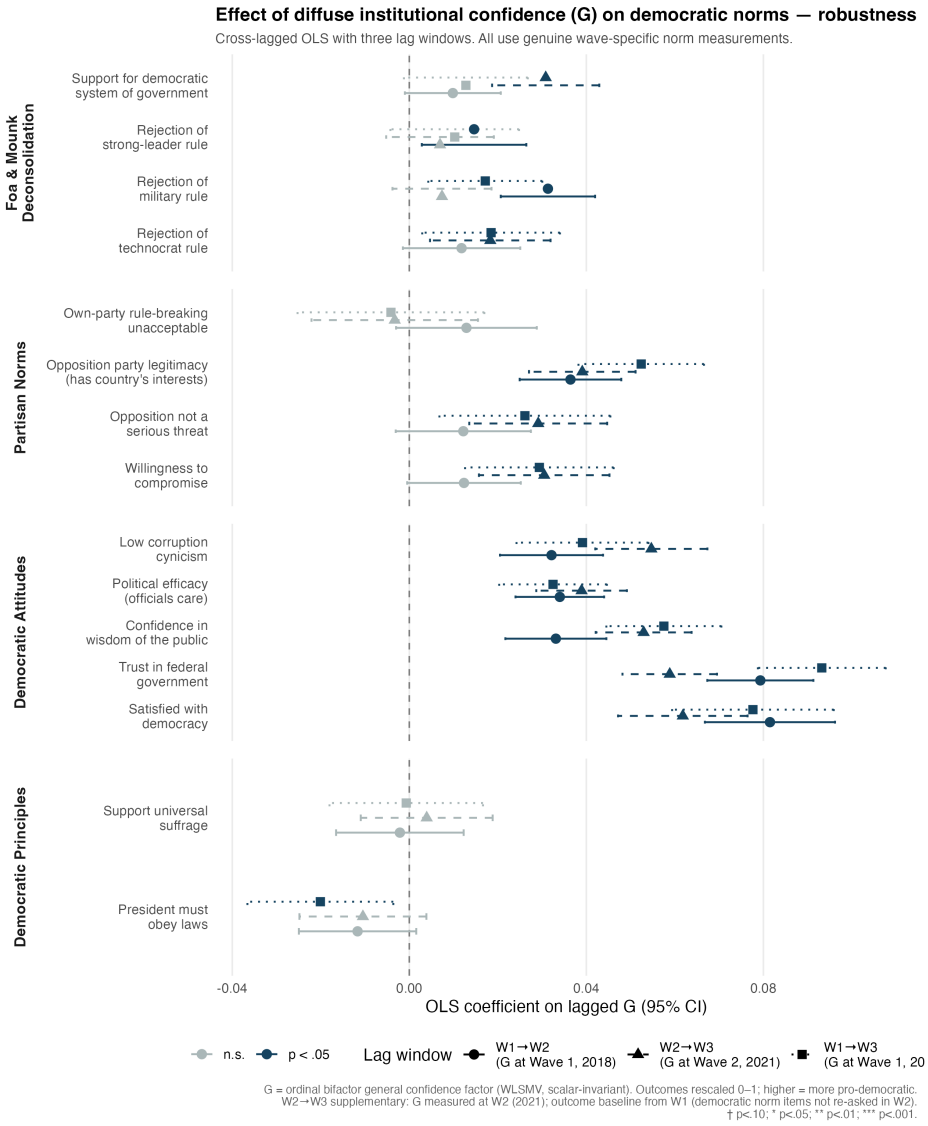

Figure 9: Effect of lagged generalized institutional confidence (G) on democratic norm outcomes across all three lag windows (Wave 1→Wave 2, Wave 2→Wave 3, and Wave 1→Wave 3). Points show OLS coefficients with 95% confidence intervals; colored points indicate $p < .05$. Controls, scaling, and estimator as in Figure 6.

Appendix H: Party-Stratified Lagged-G Effects

This appendix displays the lagged-G effects estimated separately within Democrats, Independents, and Republicans. The effects on government trust and democratic satisfaction are of nearly identical magnitude across strata.

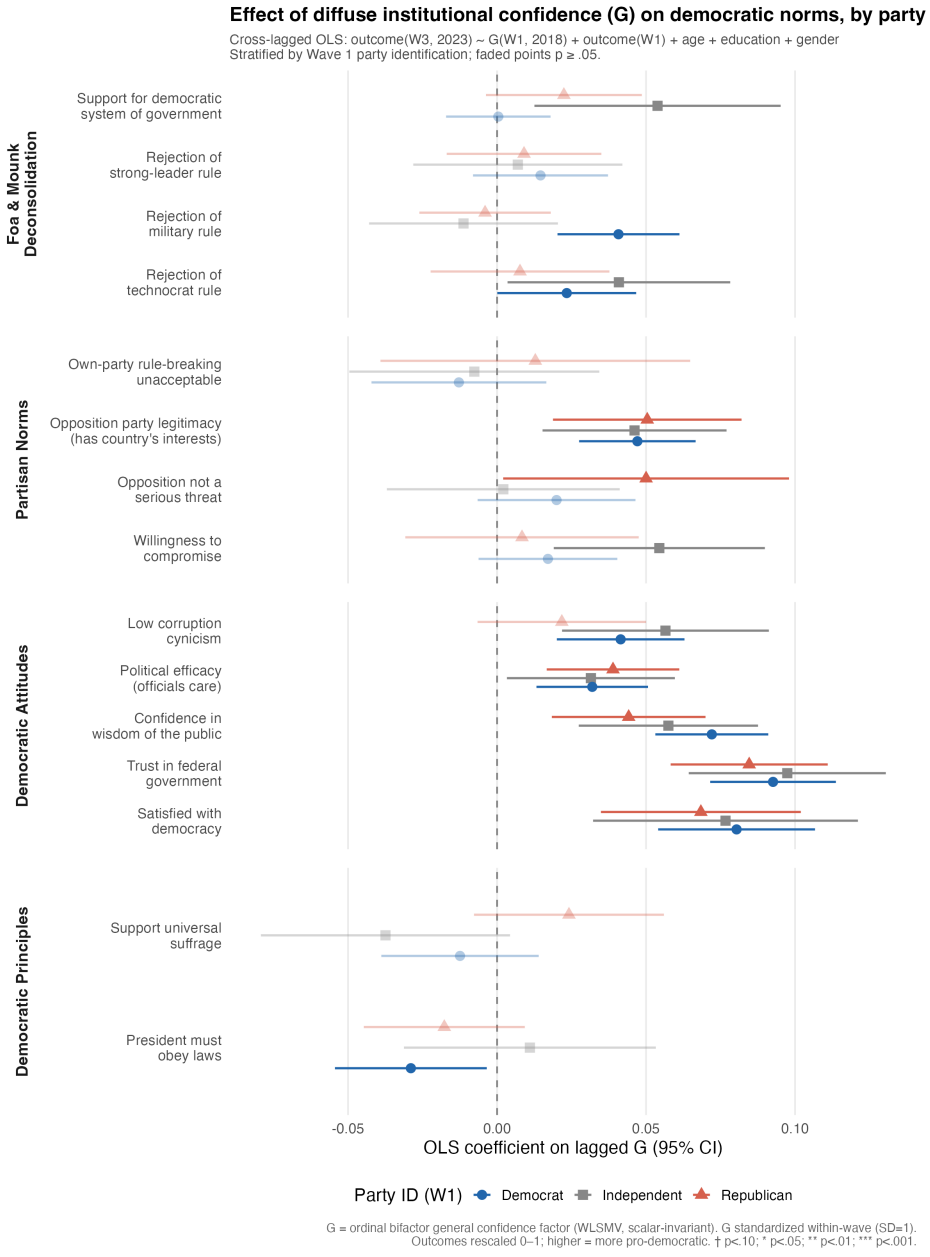

Figure 10: Party-stratified lagged-G effects on democratic norm outcomes (Wave 1→Wave 3), estimated separately for Democrats, Independents, and Republicans.

Appendix I: Within-Person Change-Score Correlations

This appendix plots the within-person change-score correlations between ΔG and change in each democratic-norm outcome over the full panel. It corroborates the cross-lagged results through a design that makes no assumption about baseline levels; the rank ordering of outcomes closely matches the OLS estimates, with the correlation for government trust reaching $r = .259$.

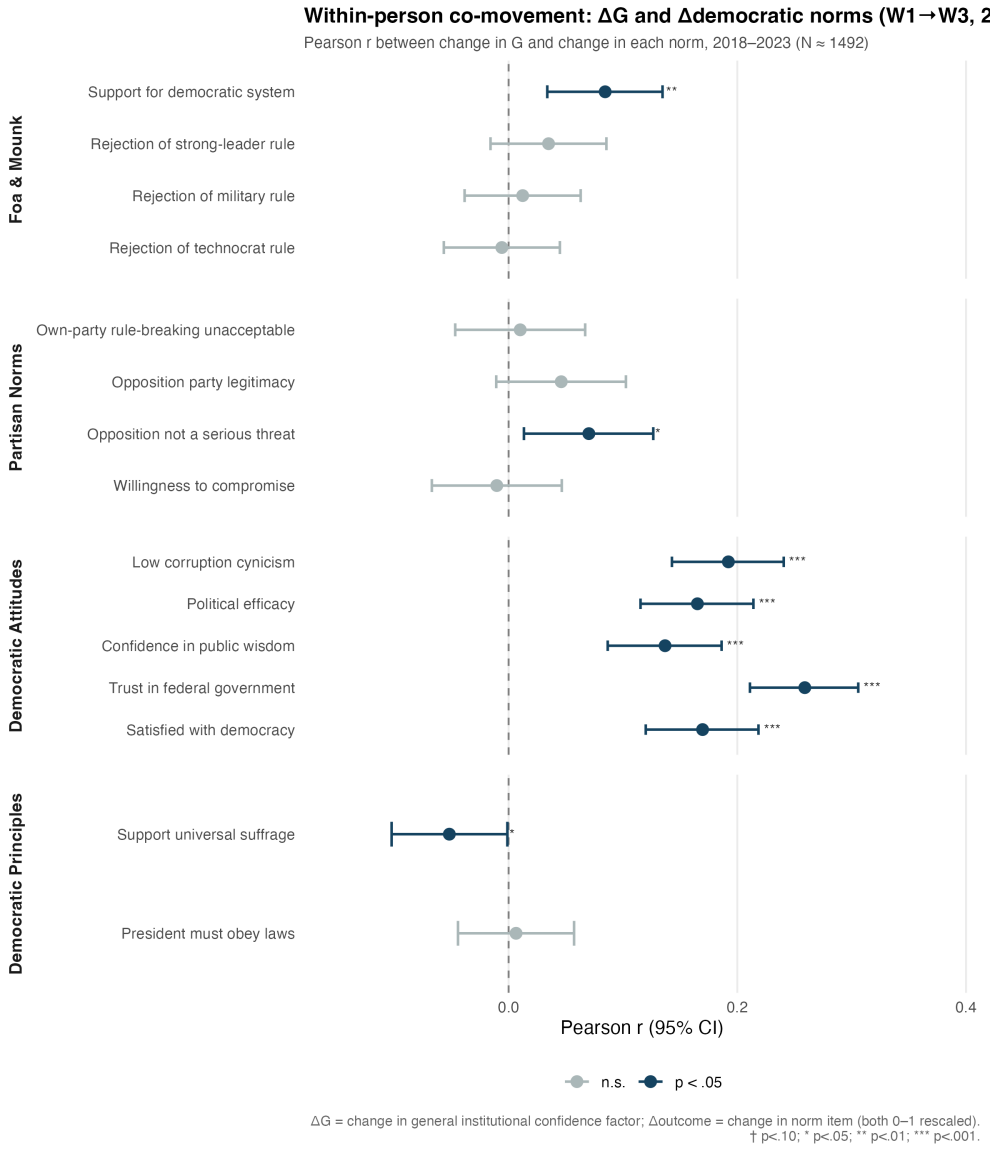

Figure 11: Within-person change-score correlations between ΔG and change in each democratic norm outcome (Wave 1→Wave 3).

Appendix J: Robustness to Survey Weighting and Attrition

Our main models are estimated unweighted. This appendix re-estimates them with survey weights to gauge whether representativeness or panel attrition affects the conclusions. For the measurement models, we re-fit the bifactor CFA with each wave's cross-sectional survey weights and re-estimate the over-time latent means; for the downstream analysis (Analysis V), we re-estimate the cross-lagged models with inverse-probability-of-retention weights, i.e. the inverse probability of remaining in the panel across the relevant waves.

The results are substantively unchanged throughout. The share of common variance carried by G rises across waves under weighting just as it does unweighted, and the domain shares decline in parallel (Table 8). The over-time decline in G is essentially the same (Table 9). And the downstream effects on democratic norms are nearly identical unweighted and retention-weighted: the largest shift for any outcome is about 0.01, and every Democratic Attitudes effect remains significant at $p < .001$ (Table 10).

Table 8: Explained common variance, unweighted vs. cross-sectionally weighted, by wave.

	ECV(G)		ECV(STATE)		ECV(PRIVATE)	
Wave	Unwt.	Wtd.	Unwt.	Wtd.	Unwt.	Wtd.
1	0.199	0.233	0.450	0.427	0.350	0.339
2	0.406	0.309	0.330	0.386	0.264	0.305
3	0.470	0.432	0.296	0.324	0.234	0.245

Notes: ECV from the ordered-categorical bifactor model (WLSMV/probit), estimated unweighted and with wave-specific cross-sectional survey weights. The rise in ECV(G) and the decline in the domain shares both hold under weighting.

Table 9: Latent G means over time, unweighted vs. cross-sectionally weighted (Wave 1 anchored at 0).

Wave	Unweighted	CS-weighted
1	0.000 (ref.)	0.000 (ref.)
2	-0.294	-0.311
3	-0.267	-0.234

Notes: Latent G means from the over-time multi-group CFA, unweighted and with cross-sectional survey weights, in Wave 1 latent SD units.

Table 10: Retention-weighted robustness for the downstream analysis (Analysis V, Wave 1→Wave 3). Effect of lagged G, estimated unweighted and with inverse-probability-of-retention weights on the same sample.

Outcome	N	Unweighted b (SE)	Retention-weighted b (SE)
Democratic attitudes			
Trust in federal government	1500	0.093*** (0.007)	0.089*** (0.007)
Satisfied with democracy	1500	0.078*** (0.009)	0.075*** (0.008)
Political efficacy	1500	0.032*** (0.006)	0.032*** (0.006)
Low corruption cynicism	1500	0.039*** (0.008)	0.037*** (0.007)
Confidence in public wisdom	1500	0.058*** (0.007)	0.056*** (0.007)
Partisan norms			
Own-party accountability	1198	-0.004 (0.011)	0.006 (0.010)
Opposition legitimacy	1198	0.052*** (0.007)	0.058*** (0.007)
Opposition not a threat	1198	0.026** (0.010)	0.016† (0.009)
Willingness to compromise	1198	0.029*** (0.009)	0.035*** (0.008)
Foa & Mounk deconsolidation			
Support democratic system	1500	0.013† (0.007)	0.020** (0.007)
Reject strong-leader rule	1500	0.010 (0.007)	0.010 (0.007)
Reject military rule	1500	0.017** (0.007)	0.017** (0.006)
Reject technocrat rule	1500	0.018* (0.008)	0.022** (0.007)
Democratic principles			
Support universal suffrage	1500	-0.001 (0.009)	0.003 (0.008)
President must obey laws	1499	-0.020* (0.008)	-0.025** (0.008)

Notes: Coefficients are the effect of Wave-1 G on the Wave-3 outcome, net of the baseline outcome, party identification, age, education, and gender. Weights are the inverse probability of Wave-1-to-Wave-3 retention. Standard errors in parentheses. $\dagger p < .10$; $^{*}p < .05$; $^{**}p < .01$; $^{***}p < .001$.