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Conventional and Unconventional Participation in Latin America: A Hierarchical Latent Class Approach*

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Building on Alvarez, Levin, and Núñez (2017), we implement a hierarchical latent class model to analyze political participation from a comparative perspective. Our methodology allows simultaneously: (i) estimating citizens' propensity to engage in conventional and unconventional modes of participation; (ii) classifying individuals into underlying "types" capturing within- and cross-country variations in participation; and (iii) assessing how this classification varies with micro- and macro-level factors. We apply our model to Latin American survey data. We show that our method outperforms alternative approaches used to study participation and derive typologies of political engagement. Substantively, we find that the distribution of participatory types

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is similar throughout the continent, and that it correlates strongly with respondents' socio-economic characteristics, ideological preferences and crime victimization.

This article proposes a hierarchical latent class modeling approach to analyze political participation in Latin America. Our work builds on Alvarez, Levin, and Núñez (2017), who develop a theoretical framework and a two-dimensional latent class model to study participation in Argentina using survey data. Here we adopt a broadly comparative perspective, expanding their analysis to 17 Latin American countries. In order to do so, we extend Alvarez, Levin, and Núñez (2017)'s empirical method, letting the parameters of the latent class model vary between countries while preserving the cross-national comparability of our results. We also adopt a more efficient estimation strategy to assess the influence of micro- and macro-level variables on political participation.

Alvarez, Levin, and Núñez (2017) distinguish between a conventional and an unconventional dimension of political participation, and examine how different activities - such as attending meetings of a political party, contacting government officials, requesting help from local authorities, protesting and striking - relate to each dimension. Rather than imposing arbitrary restrictions on the mapping between political activities and participatory dimensions (i.e., assuming a priori that certain activities are either conventional or unconventional), they use a mixture model with two categorical latent variables to estimate the strength of the association between each activity and dimension. Their empirical strategy also allows them to compute individuals' overall propensities to engage in conventional and unconventional modes of participation and, based on these propensities, to classify citizens into four participatory types, summarized in Table 1.

TABLE 1 *Dimensions and Types of Political Participation*

			Unconventional Dimension	
			Propensity	
			High	Low
Conventional Dimension	Propensity	High	*Activist*	*Conventional*
		Low	*Agitator*	*Outsider*

Note: The table summarizes the relationship between individuals' propensity to engage in conventional (rows) and unconventional (columns) forms of participation, and their participatory types (cells). Based on Alvarez, Levin, and Núñez (2017).

A citizen with a high (low) score on both dimensions is labeled an "activist" ("outsider"). Someone with a high propensity to join in conventional forms of participation but less prone to partake in activities predominantly linked to the unconventional dimension is classified as "conventional". Finally, an individual who scores highly on the unconventional dimension but exhibits little proclivity to get involved in more conventional activities is tagged as an "agitator".

The multi-level specification we propose here allows applying this framework beyond a single country, estimating individuals' underlying predispositions to engage in conventional and unconventional modes of participation for all the nations in our study and calculating the prevalence of the four participatory types throughout Latin America. Importantly, the hierarchical structure of our model accommodates potential measurement non-invariance, namely, the possibility that identical questions about political participation may have different meanings or interpretations in different settings. This has become a prominent area of study for survey methodologists in recent years (e.g. Saiegh 2015).

We fit our hierarchical latent class model using an approach that differs from that employed by Alvarez, Levin, and Núñez (2017). In their article, estimation proceeds in two sequential steps. The first stage implements a latent class model without covariates to compute survey respondents' scores on the conventional and unconventional dimensions and obtain their posterior probabilities of type assignment. To account for the fact that the types are estimated (rather than observed), a large number of types are drawn for each respondent from these posterior distributions. The second step then regresses each set of sampled types on explanatory variables. Estimates for the coefficients of these predictors are obtained by combining the output of the various second-stage models.

By contrast, we use a single-step procedure that concomitantly estimates individuals' dimension-specific propensities, their probabilities of type assignment, and the impact of micro- and macro-level factors on such probabilities. Doing so enables us to directly reflect the uncertainty in survey respondents' allocation to types and integrate it into inferences about the coefficients of the explanatory variables. This "unified" method does a better job accounting for the measurement error of classifications, yielding more accurate standard error estimates and improving the separation between classes (Kamata et al. 2018).

Besides extending the method in Alvarez, Levin, and Núñez (2017), our model also improves on previous research using cluster analysis to develop political typologies of Latin American survey respondents (e.g. Carlin 2011). Unlike our approach, cluster analysis is not based on a statistical model. Hence, it does not yield information about the probabilities of type assignment and ignores classification uncertainty, which can result in high rates of mis-classification (Kamata et al. 2018). Additionally, cluster analysis provides no straightforward way of assessing the impact of covariates on type allocation. While researchers do sometimes incorporate the cluster indicators in subsequent explanatory regression models, they typically neglect classification measurement error. This leads to biased estimates for the relationship between the classifications and the covariates of interest (Haagenars 1993). Furthermore, standard cluster analysis cannot be readily applied to cross-national data without imposing strong assumptions about the items used to derive individuals' classification – i.e., that these items are identically understood and interpreted by survey participants in all countries (De Jong, Steenkamp, and Fox 2007). Our method overcomes these limitations.

A Hierarchical Latent Class Model of Political Participation

Like Alvarez, Levin, and Núñez (2017), we simultaneously examine individuals' decisions to take part in multiple political activities in order to: (i) assess the strength of the association between each activity and participatory dimension; (ii) estimate individuals' underlying predispositions to engage in conventional and unconventional modes of participation. However, whereas in Alvarez, Levin, and Núñez (2017) the parameters linking each activity and dimension are held constant for all survey respondents, we let these relationships vary across countries through a hierarchical specification. In doing so, we build on De Jong, Steenkamp, and Fox (2007), who proposed a multi-level item response model that allows for meaningful cross-national comparisons without the need for restrictive measurement invariance assumptions. We adapt this approach and integrate it into our two-dimensional hierarchical latent class specification.

In our model, the probability that individual i, $i = 1, \ldots, N_k$, in country $k = 1, \ldots, K$, participates in activity $j = 1, \ldots, J$, is given by:

$$Y_{i,j,k} \sim Bernoulli(p_{i,j,k}) \tag{1}$$

$$p_{i,j,k} = F\Big(\alpha_{j,k} + \alpha_{c,j,k}(T_{c,i,k} - 1) + \alpha_{u,j,k}(T_{u,i,k} - 1)\Big) \tag{2}$$

where $Y_{i,j,k}$ is a binary variable based on i's response to survey item j; $T_{c,i,k}$ and $T_{u,i,k}$ are categorical latent variables taking values 1 (*low*) or 2 (*high*), denoting i's score on the conventional and unconventional dimensions of political participation, respectively; $\alpha_{c,j,k}$ and $\alpha_{u,j,k}$ are non-negative coefficients – akin to "factor loadings" – measuring the strength of the association between each participatory dimension and activity j in country k; $\alpha_{j,k}$ is an intercept; and $F(\cdot)$ is some cumulative density function - e.g., logistic or normal.

The α parameters in Equation (2) are specified as country-specific random effects:

$$\alpha_{j,k} \sim Normal(\mu_{\alpha_j}, \sigma^2_{\alpha_j}) \tag{3}$$

$$\alpha_{c,j,k} \sim Truncated\ Normal(\mu_{\alpha_{c,j}}, \sigma^2_{\alpha_{c,j}}, 0, \infty) \tag{4}$$

$$\alpha_{u,j,k} \sim Truncated\ Normal(\mu_{\alpha_{u,j}}, \sigma^2_{\alpha_{u,j}}, 0, \infty) \tag{5}$$

Equations (3)-(5) let the average probability of partaking in activity j as well as the relationship between j and the latent variables T_c and T_u differ across polities. The country-specific intercept and slopes are linked through common activity-specific means $\mu_{\alpha_j}, \mu_{\alpha_{c,j}}, \mu_{\alpha_{u,j}}$ and variances $\sigma^2_{\alpha_j}$, $\sigma^2_{\alpha_{c,j}}$, $\sigma^2_{\alpha_{u,j}}$ to ensure that T_c and T_u are measured on the same scale throughout the region (De Jong, Steenkamp, and Fox 2007).[1]

To complete the model specification, i's propensity to engage in conventional and unconventional modes of participation is expressed as a function of individual and macro-level covariates, denoted respectively by $X_{i,k}$ and Z_k:

$$P(T_{c,i,k} = high) = \frac{\exp\left(X'_{i,k}\beta_c + Z'_k\gamma_c + \eta_{c,k}\right)}{1 + \exp\left(X'_{i,k}\beta_c + Z'_k\gamma_c + \eta_{c,k}\right)} \tag{6}$$

[1]For identifiability, $\alpha_{c,j',k}$ and $\alpha_{u,j'',k}$ for a given j' and j'', $j' \neq j''$, are set to zero $\forall k$.

$$P(T_{u,i,k} = high) = \frac{\exp\left(X'_{i,k}\beta_u + Z'_k\gamma_u + \eta_{u,k}\right)}{1 + \exp\left(X'_{i,k}\beta_u + Z'_k\gamma_u + \eta_{u,k}\right)} \tag{7}$$

where $\eta_k = (\eta_{c,k}, \eta_{u,k})' \sim Normal(0, \Sigma_\eta)$ are bivariate random effects accounting for unobserved heterogeneity and intra-country correlation in the probability of scoring highly on T_c and T_u. Estimates for β, γ, η and Σ_η are obtained alongside those for α, μ and σ^2, in a single step.

We resort to Markov chain Monte Carlo (MCMC) simulations to fit the model (Lynch 2007). The Bayesian inferential framework is particularly appealing in our setting, since the number of countries under study is too small to satisfy the asymptotic criteria required by maximum likelihood estimation of multi-level models. In contrast, previous work (e.g., Gelman 2006) has demonstrated that Bayesian methods yield accurate estimates of the regression parameters and variance components of hierarchical models even with a small number of clusters, provided the number of observations per cluster is reasonably large – as is the case in our application.

Another fundamental advantage of the Bayesian approach is that we can obtain values for $T_{c,i,k}$ and $T_{u,i,k}$ at each iteration of the MCMC algorithm, drawing samples from their full conditional distributions $\{p_{c,i,k,low}, p_{c,i,k,high}\}$ and $\{p_{u,i,k,low}, p_{u,i,k,high}\}$, with:

$$p_{d,i,k,high} = \frac{P(T_{d,i,k} = high) \times \prod_j^J P(Y_{i,j,k} | T_{d,i,k} = high)}{\sum_{l=low}^{high} P(T_{d,i,k} = l) \times \prod_j^J P(Y_{i,j,k} | T_{d,i,k} = l)} \tag{8}$$

and $p_{d,i,k,low} = 1 - p_{d,i,k,high}$ for $d = c, u$.

Based on these sampled values, i is classified as an "activist" if $\{T_{c,i,k}, T_{u,i,k}\} = \{high, high\}$; as "conventional" when $\{T_{c,i,k}, T_{u,i,k}\} = \{high, low\}$; as an "agitator" if $\{T_{c,i,k}, T_{u,i,k}\} = \{low, high\}$; and as an "outsider" if $\{T_{c,i,k}, T_{u,i,k}\} = \{low, low\}$. Since i's type may vary across iterations, classification uncertainty is accounted for during the

MCMC simulations and incorporated in inferences about β, γ, η and Σ_η without the need for post-estimation routines, as in Alvarez, Levin, and Núñez (2017).

The results reported below were obtained using a standard normal cumulative distribution function in Equation (2), which allows deriving closed-form conditional distributions for $\alpha_{j,k}$, $\alpha_{c,j,k}$ and $\alpha_{u,j,k}$ and updating their values through Gibbs sampling (Lynch 2007). These conditional posteriors do not have closed forms if a logistic link is specified, but random-walk Metropolis steps can be employed to draw samples from α. While the estimates are substantively similar, execution time increases by more than 40% in this case. The Online Appendix provides additional estimation details.

Before discussing our findings, it is worth noting that the number of categories of T_c and T_u is dictated by our analytical framework, taken from Alvarez, Levin, and Núñez (2017). This is a usual practice in latent class analyses when there are theoretical expectations regarding the nature of the groups underlying the data (Oberski 2016). Nonetheless, we also estimated a factor-analytic version of our model that does not impose restrictions on the number of participatory types. As we show in the Online Appendix, the results provide empirical support for our theoretically-derived types.

Application

Data

We fit our model to data from the 2012 AmericasBarometer survey of the Latin America Public Opinion Project. The 2012 questionnaire covers a wider range of political activities for a greater number of countries than other waves, making it ideal for illustrating the application of our model. Our sample comprises 26,227 respondents from 17 countries.[2]

Our dependent variables measure respondents' engagement in the following activities: voting; attending municipal meetings; contacting municipal, local or national authorities; attending meetings of a committee of improvements; helping solve problems in the community; attending meetings of a political party; signing petitions; sharing or reading political information through social media; participating in peaceful protests; and blocking roads. For identification purposes (footnote 1), the relationships between T_c and joining roadblocks and between T_u and attending municipal meetings are restricted to zero.

Drawing on previous studies (Desposato and Norrander 2008; Carreras and Castañeda-Angarita 2014; Alvarez, Levin, and Núñez 2017), our micro-level explanatory variables are: *Age*; *Female*; *Education*; *Relative Income*, a measure of respondents' economic well-being relative to their countries' average income; *Perceived Corruption*, recording individuals' beliefs about the pervasiveness of corruption among public officials; *Crime Victimization*, an indicator for survey participants who reported having been victims of a crime in the previous year; and *Ideological Distance to Incumbent*, the spatial distance between respondents' self-placement on the left-right scale and the ideological position of the incumbent party.

The country-level covariates represent institutional and economic factors that shape citizens' incentives and opportunities for participation (van der Meer, van Deth, and Scheepers 2009; Katz and Levin 2018). These include: the degree of respect for the *Rule of Law*; the presence and enforcement of *Compulsory Voting* laws; the effective number

[2]Table A.1 in the Online Appendix provides sample sizes for each country. Section A.3 reports estimates from additional specifications covering a larger number of countries and years and using alternative operationalizations for the explanatory variables. The main findings remain unchanged.

of parties (*ENPP*); *GDP per capita*; and *Social Spending* as a percentage of GDP. The Online Appendix provides supplementary information about variable definitions, coding and sources.

Results

Table 2 reports posterior summaries for the parameters capturing the association between each political activity and participatory dimension, averaged across countries.

The estimates are generally consistent with expectations. Activities considered as conventional by the literature, like attending municipal meetings and contacting the authorities, have the largest values of α_c, while α_u is largest for activities deemed unconventional, such as protesting and joining roadblocks (van der Meer, van Deth, and Scheepers 2009). The differences in the relative strength of the association between these activities and each participatory dimension are statistically significant: α_c is more than two standard deviations larger than α_u for *Contacting Municipality* and *Contacting Local Authority*, while the opposite holds for *Protesting*. Each of these activities is thus predominantly related to a single dimension.

Other activities like petitioning or attending party meetings exhibit relatively large values of both α_c and α_u, implying that they reflect conventional and unconventional latent predispositions towards political action. More generally, the values of α_c and α_u are statistically indistinguishable for the majority of the activities, meaning that they defy straightforward binary classifications. Their "dual nature" would be missed by ad-hoc typologies assuming a priori that certain forms of participation are either conventional or unconventional. By contrast, our method allows – but does not force – each item to be related to both dimensions, yielding a data-driven classification of political activities.

TABLE 2 *Posterior summaries for $\alpha_{c,j,k}$ and $\alpha_{u,j,k}$ averaged across countries*

Activity	**Dimension**	
	(1) Conventional	(2) Unconventional
Voting	0.451 (0.051, 0.869)	0.217 (0.004, 0.744)
Municipal meeting	1.308 (0.935, 1.652)	0.000 (0.000, 0.000)
Contacting municipality	1.701 (1.183, 2.301)	0.355 (0.020,0.811)
Contacting local authority	1.632 (1.015, 2.323)	0.259 (0.009, 0.792)
Contacting national authority	1.200 (0.694, 1.760)	0.361 (0.028, 0.793)
Improvements meeting	1.322 (0.736, 1.777)	0.590 (0.106, 1.258)
Solving community problems	1.126 (0.573, 1.563)	0.571 (0.109, 1.240)
Party meeting	1.009 (0.549, 1.393)	0.774 (0.230, 1.383)
Petitioning	0.949 (0.065, 1.433)	1.173 (0.586, 1.909)
Sharing online	0.300 (0.018, 0.677)	0.779 (0.141, 1.559)
Protesting	0.367 (0.024, 0.796)	2.924 (1.468, 4.123)
Blocking	0.000 (0.000, 0.000)	2.331 (1.242, 3.449)

Note: The table reports posterior means and 95% credible intervals (in parentheses) for $\alpha_{c,j,k}$ and $\alpha_{u,j,k}$ averaged across k.

Figure A.1 in the Online Appendix reports estimates of $\alpha_{c,j,k}$ and $\alpha_{u,j,k}$ for each country. The general patterns are consistent with those in Table 2: despite cultural, political and socio-economic differences between Latin American democracies, the relationship between political activities and participatory dimensions is similar throughout the continent. However, the figure also reveals some cross-national differences in the magnitude of the estimates and in the strength of the association between activities and dimensions. These differences highlight the importance of modeling $\alpha_{c,j,k}$ and $\alpha_{u,j,k}$ as country-specific random effects.[3]

In this direction, Table 3 shows that our baseline model (column 1) outperforms more restrictive specifications assuming equality of activity-specific slopes (column 2) or slopes and intercepts (column 3) across countries. Relaxing these invariance constraints improves fit according to a variety of commonly used model selection criteria.[4]

Column (4), in turn, reports goodness-of-fit statistics for a model allowing the α parameters to vary across countries but assuming that the latent propensities to engage in conventional and unconventional participation are the same for all respondents. This specification thus assumes that all respondents belong to the same participatory type. A comparison between columns (1) and (4) reveals that our baseline specification is favored by every model selection criterion. Hence, accounting for heterogeneity in political engagement in terms of a small yet substantively meaningful number of types greatly improves model fit.

[3]Figure A.2 in the Online Appendix also shows cross-national variations in the intercepts $\alpha_{j,k}$.

[4]See Ntzoufras (2011) and Gelman, Wang, and Vehtari (2014) for a description of the formulas used to calculate these measures.

TABLE 3 *Comparing our model vis-á-vis alternative specifications*

	(1) Baseline model	(2) Model assuming metric invariance	(3) Model assuming scalar invariance	(4) Model assuming a single type
Akaike Information Criterion (AIC)	207,501.97	227,298.54	277,770.95	229,600.25
Bayesian Information Criterion (BIC)	212,504.80	229,162.34	278,065.23	234,586.72
Consistent AIC (CAIC)	213,116.80	229,390.34	278,101.23	235,196.72
Deviance Information Criterion (DIC)	207,269.23	227,248.27	277,707.77	240,189.24
Watanabe-Akaike Criterion (WAIC)	207,279.37	227,244.46	277,708.72	240.805.32
χ^2 (p-value)		<0.001	<0.001	<0.001

Note: Column (1) is our preferred specification. Column (2) assumes that $\alpha_{c,j,k}$ and $\alpha_{u,j,k}$ are invariant across countries. Column (3) additionally assumes that $\alpha_{j,k}$ are the same $\forall\ k$. Column (4) allows $\alpha_{j,k}$, $\alpha_{c,j,k}$ and $\alpha_{u,j,k}$ to vary across countries but assumes that $T_{c,i,k} = T_{u,i,k} = 2\ \forall\ i$. For AIC/BIC/CAIC/DIC/WAIC, differences larger than 10 provide overwhelming evidence in favor of the model with the lower value (Ntzoufras 2011) – which in all cases is the model in column (1). The χ^2 tests compare the fit of the most parsimonious specifications - columns (2), (3) and (4) - against column (1); again, the p-values indicate that our preferred model fits the data significantly better.

Figure 1 provides information about the relative prevalence of the four types in our sample. As noted before, our method does not treat respondents' types as fixed, but instead estimates the probability that each individual is assigned a high conventional and high unconventional type. The upper panel of the figure plots $P(T_{c,i,k} = high)$ and $P(T_{u,i,k} = high) \, \forall \, i$. On average, the probabilities that a survey respondent scores highly on the conventional and unconventional dimensions are 0.22 and 0.07, respectively. Hence, in line with prior findings (Klesner 2007), our results indicate that Latin Americans' propensity to engage in politics, even through relatively less disruptive or more routinized means, is rather low. Based on these dimension-specific propensities, the mean posterior probabilities of classifying respondents into each of the types are: 0.73 for "outsiders", 0.20 for "conventionals", 0.05 for "agitators", and 0.02 for "activists".

The posterior probabilities of type assignment are very precisely estimated, as seen in the lower panel of the figure, which displays their 95% credible intervals. The lack of overlap between the intervals indicates that the types are well separated (see also Figures A.3 and A.4 in the Online Appendix). This implies that the different types capture distinct patterns of political participation present in our sample, and that respondents can be accurately and unambiguously assigned to a particular type (Depaoli 2013). As we noted in the introductory section, the ability to improve class separation is one of the advantages of our "unified" estimation procedure.

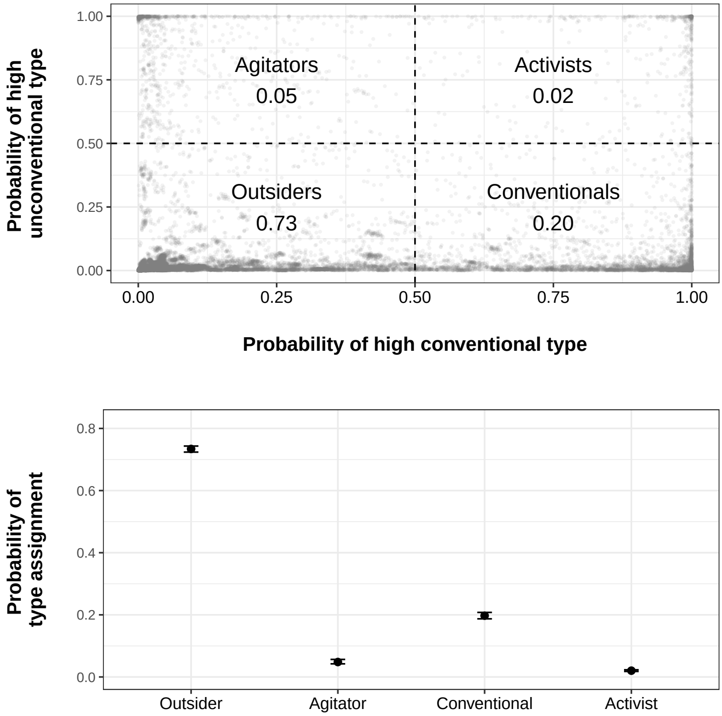

Figure 1. *Probabilities of type assignment*
Note: The upper panel depicts the relationship between the probability of being assigned a high conventional and high unconventional type. Circles represent survey respondents. The lower panel reports posterior means and 95% credible intervals for the probabilities of type assignment.

Moreover, the probabilities of type assignment are statistically indistinguishable across countries, as seen in Figure 2.[5] This is a remarkable finding, as our model did not impose any restrictions on the distribution of types across Latin America. There are, on the other hand, noticeable within-country differences in the probabilities of type assignment. Survey participants are significantly and decidedly (between 1.5 and 2 times) more likely to be

[5]We show in the Online Appendix (Tables A.4 and A.5) that the latent classes are also well separated in each nation.

outsiders than to be allocated to the other three categories in each and every one of the democracies under study. At the other extreme, respondents throughout the region are least likely to belong to the activist type.

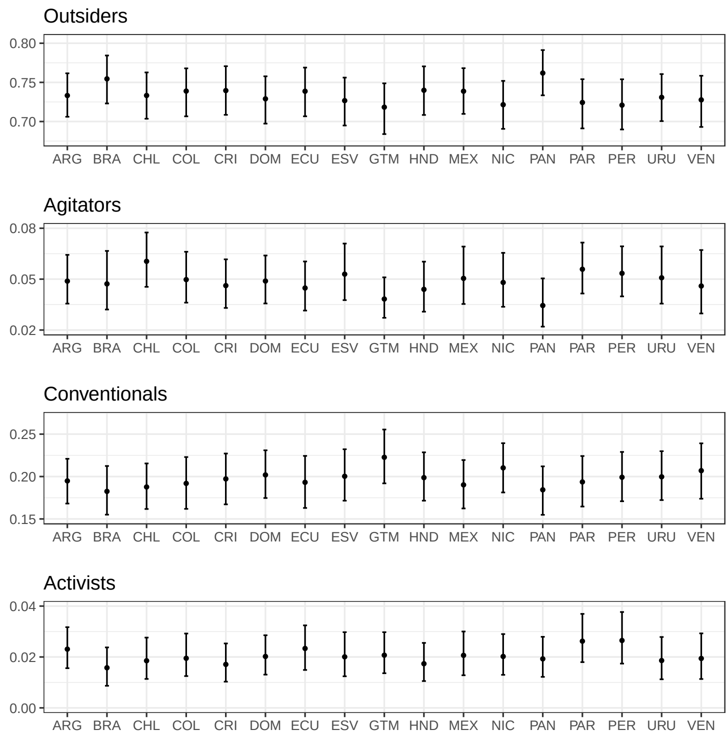

Figure 2. Probabilities of type assignment, by country
Note: Circles represent posterior means. Vertical lines give 95% credible intervals.

Figure 3 examines the impact of individual and contextual factors on these probabilities. Among the individual-level covariates, *Crime Victimization* is strongly and significantly correlated with type allocation. In consonance with prior research (e.g., Bateson 2012), we find that being victim of a crime is associated with an increase in political engagement:

survey participants who reported recent crime victimization are 9.18 percentage points less likely to be outsiders than comparable non-victims. The estimates for socio-demographic

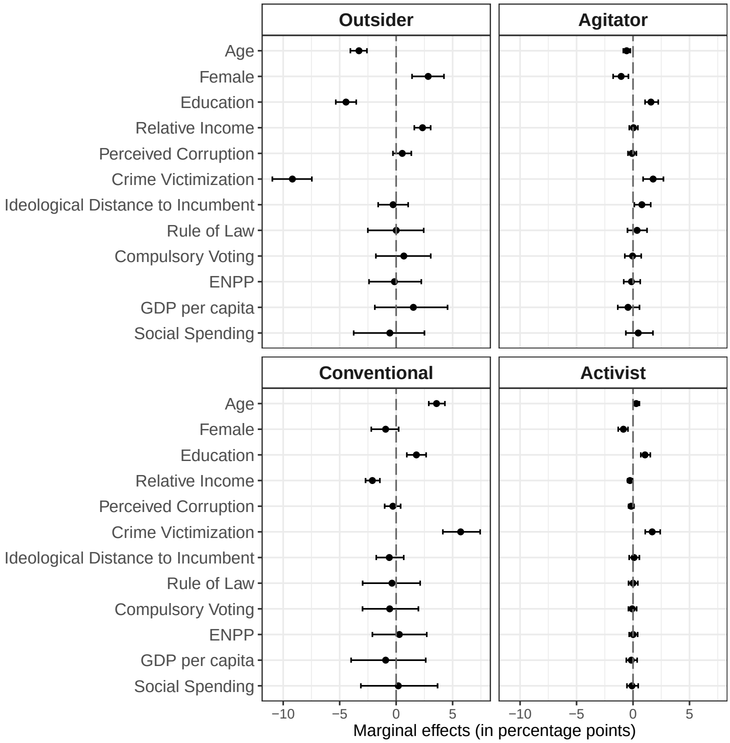

Figure 3. Expected change in the probabilities of type assignment associated with a (unit) change in the covariates
Note: Circles represent posterior means. Horizontal lines give 95% credible intervals.

characteristics are also consistent with the literature (Desposato and Norrander 2008; Carreras and Castañeda-Angarita 2014): older, male and more educated respondents are less likely to be classified as outsiders and more prone to engage in both conventional and unconventional forms of participation than younger, female and less educated individuals.

Particularly interesting are the estimates for *Ideological Distance to Incumbent*. Each one-unit increase in this variable is associated with a 0.78 percentage point increase in the probability of being classified as an agitator; its impact on the other type assignment probabilities is statistically insignificant. Hence, instead of “withdrawing” from politics or engaging in conventional forms of activism, citizens less ideologically aligned with the incumbent party turn to unconventional forms of participation.

On the other hand, none of the country-level covariates has a direct impact on individuals’ allocation to participatory types (see also Figure A.5 in the Online Appendix). Nonetheless, contextual factors might indirectly shape allocation to types through the incentives and information they provide to citizens. For instance, prior work (Katz and Levin 2018 and Dassonneville and McAllister 2020, among others) has underscored the influence that *ENPP* and *Compulsory Voting* exert on individuals’ attitudes, behavioral incentives and political sophistication, which may affect their decision to engage in conventional and/or unconventional forms of participation. In this sense, Table A.6 in the Online Appendix shows that adding country-level covariates improves the explanatory power of our model even if none of these variables is statistically significant in isolation.

For comparison, Figure A.7 in the Online Appendix reports covariate “marginal effects” estimated from a model that ignores classification uncertainty, as is the standard practice in cluster analysis. In line with the arguments of Haagenars (1993) and Kamata et al. (2018), the figure highlights that neglecting classification measurement error leads to different estimates - and, in some cases, different conclusions - regarding the influence of individual and contextual variables on political participation in Latin America.

Conclusions

This research note contributes to the comparative study of political participation by simultaneously examining individuals' propensity to engage in conventional and unconventional activities across 17 Latin American nations. We draw on Alvarez, Levin, and Núñez (2017), extending their latent class model to the cross-national setting and implementing an estimation approach that is better able to account for classification uncertainty and integrate it into inferences about the model parameters.

Our data reveal that, in spite of individual and contextual differences, Latin Americans combine the different activities in a similar fashion. At the same time, we show that accommodating cross-national differences in the relationship between activities and participatory dimensions through a multi-level specification greatly improves the model's fit to the data. Our results also indicate that the average propensity to engage in politics is quite low across the continent: survey respondents are more likely to shun politics than to partake in either conventional or unconventional activities. Despite this predominance of outsiders, allowing for additional citizen types enhances the model's explanatory power. Our model's ability to account for within- and between-country variations in participation is especially important in view of recent political developments in Latin America. Whereas countries like Chile, Ecuador and Peru have witnessed massive protests in the last year, political discontent in other nations (e.g., Argentina, Uruguay) has been channeled through conventional electoral means. The release of public opinion surveys covering these events will provide us with the opportunity to examine how the distribution of participatory types has evolved in the region.

Although our research focused on participation, our multilevel latent class model can be adapted to analyze other political behaviors and attitudes from a comparative standpoint.

Latent class analysis has found a growing number of uses in political science, but most applications have been single-country studies. In the few instances in which these models have been fitted to multi-country data, concerns about measurement invariance have been typically ignored (e.g., Blaydes and Linzer 2008). Our hierarchical specification allows applying latent class models to cross-national surveys without the need to rely on stringent parameter restrictions or to impose unrealistic assumptions about the meaning of the items across different populations.

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