Political Divisions in Large Cities:

The Socio-Spatial Basis of Legislative Behavior in Chicago and Toronto

Zack Taylor, University of Western Ontario, zack.taylor@uwo.ca

David A. Armstrong II, University of Western Ontario, dave.armstrong@uwo.ca

ABSTRACT: Contemporary cities are frequently characterized as divided by race and socio-economic status, yet the political effects of segregation and stratification are rarely fully explored. Urban politics scholars have disagreed on whether urban politics is essentially consensual, conflicts are issue-based and transitory, or social and economic divides generate enduring political cleavages. We contribute to these debates with an analysis of elite conflict as manifested in recorded city council votes in two large, heterogeneous North American cities, Chicago and Toronto, over a multi-decade period. The analysis employs a new technique for analyzing the dimensionality of roll-call votes. We find evidence of durable coordination among ward councilors in both cities, however the substance of conflict differs. Correlating the dimensions of voting behavior with ward characteristics indicates that Chicago's aldermen divide on racial lines, while Toronto's councilors primarily divide on the place characteristics of wards, and secondarily on socio-economic status.

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<A>Introduction

Contemporary cities are frequently characterized as divided or polarized (Castells and Mollenkopf 1991; Fainstein, Gordon, and Harloe 1992; Marcuse and Van Kempen 2000). There is a rich international literature on the causes and measurement of residential segregation by race and class (Massey and Denton 1993; Logan 2013; Nightingale 2012; Trounstine 2018). Particularly in the United States, observers have focused on social and economic divisions between central cities and suburbs within metropolitan areas, fueled by domestic and international migration, globalization, deindustrialization, and the rise of the knowledge economy (Ellen 1999; Rusk 2013; Hamilton 2014). While these analyses have focused on the causes of urban socio-economic division and stratification, they rarely systematically consider their electoral and political effects. In parallel, scholars of urban political economy and local politics have theorized and analyzed municipal politics, typically in central cities, from a variety of perspectives, including pluralism (Dahl 1961), elite dominance and growth coalitions (Gendron and Domhoff 2009; Hunter 1953; Stone 1989), Marxian political economy (Logan and Molotch 2007), and cooperation and conflict between racial groups (Colburn and Adler 2005; Benjamin 2017). These perspectives generate different expectations regarding the scope of urban politics and the bases and durability of urban political conflicts. They also differ in terms of methods used and information analyzed, including qualitative case studies and quantitative analyses of individual-level surveys and aggregate election results.

We contribute to this literature with an analysis of elite behavior as manifested in recorded city council votes. We do so in two large North American cities, Chicago and Toronto, over multiple terms of office spanning multiple decades – 1981–2019 in Chicago (nine terms over 38 years) and 1998–2018 in Toronto (seven terms over 20 years).[1] Legislative decisions are an underexploited resource for analyzing the nature of political conflict in cities (Hajnal 2014, 527). The spatial modelling of roll-call votes locates each legislator in relation to every other legislator on one or more dimensions based on their recorded approval of or opposition to legislative proposals. Through this, the complexity of hundreds, even thousands, of voting events over often considerable periods of time is reduced to a set of scores on a limited number of latent variables, and the clustering of similar-behaving legislators is revealed. We then uncover the meaning of the dimensions by correlating councilor ideal points with ward characteristics. For Chicago, we also investigate associations between these dimensions and councillor race. To perform this analysis we constructed longitudinal databases of councilor votes in both cities by digitizing all available printed council minutes and developed a novel procedure to analyze them. We also constructed original databases of the characteristics of council members and their wards. Our procedure and databases are important contributions in their own right.

We arrive at two main findings. First, councilor voting is enduringly low-dimensional in both cities. Indeed, most of the variation in councilor voting can be reduced to two latent variables in both cities. Most councilors' ideal points are also stable across time. Low-dimensionality and positional stability suggest that legislative behavior is organized around enduring divisions despite there being few institutional constraints on legislators' choices.

Our second finding is these divisions have a socio-spatial basis. That is, councilors' dimension scores correlate with recognizable complexes of ward characteristics. Moreover, the divisions these complexes represent are very different in the two cities. For much of the period studied, legislative behavior in Chicago is essentially unidimensional and aligns with wards'

[1] Council members are called aldermen in Chicago and councilors in Toronto. We use the term "councilor" to refer to elected ward representatives in both cities.

racial composition. The first dimension captures a distinction between wards that are disproportionately home to African Americans, transit commuters, and the unemployed, and wards that are disproportionately inhabited by "ethnic" non-Hispanic whites, homeowners, and car commuters – in essence, the nexus of race and place. A second, weaker dimension captures the nexus of race and socio-economic status, pitting representatives of wards that are disproportionately home to African Americans or the unemployed against wards with high proportions of non-Hispanic whites, the university-educated, and high-cost housing. We also find that African American and white aldermen are polarized, especially on the first dimension, although this has diminished over time. In Toronto, however, the dominant dimension corresponds to a distinction between two geographic zones of the city: the walkable core and the automobile-oriented residential periphery. We propose that these two built environments give rise to material interests and social identities that carry considerable weight in Toronto's electoral politics. The second, weaker, dimension divides wards based on markers of socio-economic status: religion, educational attainment, ethnicity, and occupation.

We conclude with discussion of these findings' implications for the study of urban politics and potential extensions.

<A>Literature Review

For decades, scholars of local politics and urban political economy have used diverse methods to analyze the nature, scope, and durability of political conflicts in cities, including at the elite level. This body of work can be divided into three general categories, which we characterize as the pluralist, structuralist, and group conflict perspectives. These perspectives generate divergent expectations regarding the substance and durability of political divisions in cities generally, and the behavior of municipal legislators in particular.

The pluralist perspective portrays local government as a venue in which society's many interests are aggregated on an issue-by-issue basis (Dahl 1961). In this view, there are no durable cleavages in local politics; rather, society is composed of a multiplicity of crosscutting interests and identities that are more or less salient to any given issue. In this vein, DeLeon (1992, 13) characterizes San Francisco as "hyperpluralist" – a place with no majorities, where "everything is *pluribus*, nothing is *unum*." From this perspective we would expect there to be no enduring factions on municipal councils as each issue, or issue space, is expected to generate its own transitory sets of legislative coalitions.

Structuralist perspectives see local politics as inherently limited by contextual constraints. Drawing on public choice economics, Peterson (1981) argues that inter-municipal competition for residents and investment leads to policy convergence. Similarly, Tiebout (1956) proposed that geographic sorting of residents in search of similar policies ("voting with one's feet") would lead to homogeneous policy preferences within local government. Building on elite theory (Hunter 1953) and studies of business dominance and boosterism (Kantor and David 1988), Logan and Molotch (2007) show from a Marxian perspective how a discourse of "value-free development" – that growth-oriented urban development projects have no downside – organizes conflict out of local politics. The imperatives of market competition therefore suppress political conflict. From this perspective we would expect relative consensus among municipal legislators and frequent unanimity.

Others see local government as a venue for alliance and opposition between societal groups. Indeed, a dominant tradition in American urban politics research focuses on coalitions and conflicts between groups defined by race – African-Americans, Latinx residents, and whites (Sonenshein and Pinkus 2005; Sugrue 2005; Benjamin 2017; Trounstine 2018; Hajnal and

Trounstine 2014; Kaufmann 2010) – although the unity of such groups may sometimes be crosscut by class divisions (Stone 1989; Swanstrom 1985; Dawson 1994). Doering, Silver, and Taylor (2021) find that neighborhood-scale vote shares for mayoral candidates in Chicago, Toronto, and London are predicted by three latent variables that are strongly correlated not only with the racial and socio-economic composition of neighborhoods, but also indicators of core and suburban lifestyles, including housing types and commuting behavior. Inside and outside the United States, observers have also identified other durable group-based cleavages in local politics, including religion (Boal 2008), sexuality (Castells 1983), and language (Levine 2010; Sancton 1985).

This group representation perspective predicts that stable legislative factions will mirror enduring social and economic divisions among the population because legislators are responsive to the publics they represent. There is growing evidence that this is the case. Researchers have found that taxation and spending levels and other policy outcomes correlate with the aggregate preferences or ideological leaning of municipal residents (Einstein and Kogan 2015; Tausanovitch and Warshaw 2014; Connolly and Mason 2016). Moreover, survey-based research indicates that voters' ideological self-placement predicts vote choice in local elections (McGregor, Moore, and Stephenson 2016) and that the ideological orientation of nonpartisan local councilors is consistent with that of their constituents (Lucas 2020). While this literature suggests that councillors are responsive, its analytic focus on economic outcomes such as taxation and spending levels, and relatedly the fiscal conservatism or liberalism of the representatives and the represented, does not capture the representation of non-economic group identities based on race, ethnicity, or religion, or policy interests generated by the selective consumption of municipal services and infrastructure.

To investigate these different expectations, we make use of a seldom-used data source, municipal council roll-call votes (Bucchianeri 2020; but see Simpson 2001; Winn and McMenemy 1973). Building on foundational work by Key (1949) and Poole and Rosenthal (1997), political scientists have analyzed roll-call votes to reveal more or less durable patterns of coalition and conflict in the U.S. Congress. Stable legislative factions are interpreted as evidence of durable societal cleavages. Scholars have found that legislators' choices can in most circumstances be predicted by a single dimension, which is typically characterized as ideology (the left-right cleavage), although on some issues an orthogonal dimension emerges such as sectionalism, whereby Democratic legislators from the North behaved more like Republicans than Southern Democrats in the century after Reconstruction (Poole 2005, 140, 204). Even in contexts where a single dimension is dominant, the two-dimensional solution is often superior even if the second dimension lacks as clear a substantive interpretation as the first dimension (Armstrong et al. 2021).

Political parties play a strong coordinating role in partisan legislatures, although this effect is stronger in parliamentary systems, which feature strong party discipline, than in the American Congress, where party factions have historically reached across the aisle to form issue-by-issue coalitions (Poole 2005; Shor and McCarty 2011; Poole and Rosenthal 1997; Godbout 2020). While highly institutionalized and partisan national legislatures have received the most attention, scholars have also analyzed legislator behavior in less structured bodies. Legislators are not entirely free-floating in these contexts. In fact, they have been shown to coordinate and form stable coalitions in such contexts as state legislatures with officially nonpartisan election (Welch and Carlson 1973; Masket and Shor 2014; Wright and Schaffner 2002) and legislatures dominated by a single party (Harmel and Hamm 1986; Key 1949). Martin and Quinn (2002) have applied similar choice models to understand the behavior of U.S. Supreme Court justices, who, while appointed by the president and confirmed by the Senate, are nominally nonpartisan.

Municipal legislatures pose important conceptual questions and methodological challenges in roll-call research. Local councils in the United States and Canada are relatively unstructured; most are nonpartisan or single-party dominant, and they lack the systems of seniority and committee organization found at the national and state levels. The limited research on municipal roll-call votes lends some support to the notion that coalitions are loose and transitory. Bucchianieri's (2020, 1056) analysis of 151 partisan and nonpartisan municipal councils in the United States, finds that "voting behavior in most nonpartisan contexts is relatively disorganized" and concludes that "partisan elections – and the party organizations that nearly always come with them – are critical to structuring elite competition, such that their absence in nonpartisan governments impedes long-term cooperation." In his analysis of roll-call votes in Chicago in selected years between 1863 and 2001, however, Simpson (2001, ch. 10) finds that, more often than not, the mayor exerted strong control over aldermanic voting, although there have been periods of factional conflict.

<A>Chicago and Toronto compared

Chicago and Toronto are rich laboratories of urban politics. They were selected because they share physical, economic, and institutional features while also differing in important respects. **(See Table 1.)** Both are English-speaking cities in North America whose development occurs in the context of a market economy. Each emerged in the 19th century as manufacturing and logistics hubs on the Great Lakes but have transitioned to postindustrial service economies in recent decades. Both are central cities that anchor greater metropolitan areas, yet are themselves territorially expansive (~230 sq.mi.). Council representation is ward-based. The two cities have large and ethno-racially diverse populations that have been recently transformed by domestic and international migration. Ethno-racial groups tend to be spatially segregated in both cities (Darden and Fong 2012; Glaeser and Vigdor 2012). Reflecting postindustrialization, neighborhoods and the electoral wards they comprise are also segregated by income, occupation, and educational attainment, which in Chicago are imbricated with race (Walks 2020; Smith, Sonmez, and Zettel 2021). Each city is fully built out within its boundaries, and so urban growth occurs entirely through infill and redevelopment, fueling gentrification and other neighborhood change processes. The urban landscape of each city encompasses a diverse range of neighborhood environments. While downtown wards are dense and feature high levels of commuting by transit, walking, or cycling, more peripheral wards are on average lower density and more car-dependent. Ward office may therefore become a vehicle for the representation of group interests, be they rooted in ethnic identity or socioeconomic status, or also place-based interests associated with distinct built environments and mobility behaviors.

The two cities differ, however, in several respects. Institutionally, Chicago's outer boundaries have remained substantially unchanged since the 1930s. Chicago's mayor sits apart from council and possesses stronger prerogatives, including the ability to veto council decisions and hire and fire senior staff. The Democratic Party has long dominated Chicago politics, and so elite competition has generally been between party factions. Mayors have historically used their control over party resources to reward supporters and punish opponents, although the "machine" has weakened over time. Prior to 1999, Chicago's mayoral elections were partisan with candidates selected in primary elections. Since then, mayoral elections have been nonpartisan. If no candidate receives an absolute majority of votes cast, the top two challengers face each other in a runoff. Aldermanic elections have always used nonpartisan ballots.

Table 1: City characteristics

	City of Chicago	City of Toronto
Population	2,746,388	2,794,356
Largest ethno-racial groups	33% Non-Hispanic white 30% Black 29% Hispanic	49% white 13% South Asian 11% Chinese 9% Black 6% Filipino
% Foreign-born	21%	49%
% Age 25+ with > BA	40%	44%
Land area	590 km^2	630 km^2
Size of council	50 single-member wards	28 two-member wards (1998–2000 term) 44 single-member wards (2000–18 terms)
Partisanism	Mayoral elections were partisan until 1999, after which they became nonpartisan	All elections are nonpartisan
Mayoral powers	Does not sit as part of council; may veto council votes; administration reports to mayor	Sits on council; no veto; administration reports to council

Note: Data are from the 2020 Census (population and race) or 2015–19 ACS (other variables) in Chicago and the 2021 (total population) or 2016 Census (other variables) in Toronto.

The City of Toronto is a newer entity, the product of the 1998 amalgamation of a two-tier government that had existed since 1954. Seven general elections have occurred since amalgamation. Toronto has a "weak mayor" system in which the mayor has no veto power, votes as a member of council, and mobilizes support through persuasion. While Toronto generally supports parties of the center-left in national and provincial elections, and local politicians are often identified with provincial or federal parties, offices are officially nonpartisan and local politicians have no control over national or provincial party machinery.

Table 1 also reveals differences in ethno-racial composition. In 1970, Chicago was 60% white and 32% African American; today, the population is split almost equally between African American, non-Hispanic white, and Latinx residents. According to the 2020 Census, 21% of the city's population was foreign-born, of which half originate in Latin America. Toronto has become more diverse over time. In 1981, 41% of residents were foreign-born, two-thirds originating in the United Kingdom and continental Europe. In 1996, the first year in which "visible minority" identity was included in the census questionnaire, 37% of residents reported being non-white, of which 22% identified as Black, 22% South Asian, and 25% Chinese. By 2018, the city's population was 51% non-white and 51% foreign-born. No single ethno-racial group dominates the non-white population: 13% are South Asian, 11% Chinese, and 9% Black.

<A>Methods and Data

<B>Roll-call votes

We approached the analysis of roll-call votes without preconceived notions regarding the dimensionality of legislators' behavior and its stability over time. We developed a novel procedure, fully described in the supplementary appendix, that imposes few constraints on the analysis. The main model used is the Bayesian dynamic item response (D-IRT) model popularized by Martin and Quinn (2002) in their analysis of U.S. Supreme Court justice ideology over time. Unlike their unidimensional approach, however, we are interested in estimating models in multiple dimensions. We also seek to avoid making arbitrary decisions about which dimensions influence which items. To avoid this dilemma, Potthoff (2018) proposes using principal components analysis (PCA) to estimate the ideal points of legislators and to then use those to estimate bill parameters if they are also desired. He describes a method for allowing evolution over time in the ideal point locations of legislators but constrains time trends to be linear. We have no reason to believe that the evolution of councilor ideal points will even be monotonic, much less linear over time and thus this solution is not sufficiently flexible for our purposes. Using Potthoff's suggestion as a starting point, we adopt a multi-stage process, the technical details of which are described in the supplementary appendix. We begin by using a generalized low rank model (GLRM), which is more permissive of missing data than PCA

(Boehmke and Greenwell 2020; Udell et al. 2016), to estimate the dimensionality of all voting events. Second, we identify the votes best predicted by each of the latent variables produced. Third, we run a unidimensional Bayesian D-IRT model on each of these subsets of votes. To anchor the scores, we identify councilors who voted on at least 50 proposals together and agreed the least in their cast votes. Finally, we orthogonalize the latent dimensions through residualization. This is necessary because each D-IRT model was estimated separately.

This procedure affords several advantages over available alternatives, including Martin and Quinn's unidimensional IRT model and NOMINATE-based approaches (Armstrong et al. 2021) that assume fixed placements across time (W-NOMINATE), parametric evolution of ideal-points over time (DW-NOMINATE), or term-by-term W-NOMINATE models that allow evolution of placements but sacrifice comparability of scores between terms. First, our approach locates legislators in multidimensional space on a common scale across time. Second, each dimension is orthogonal to the others, meaning that it potentially represents a distinct axis of disagreement. Third, each legislator's position is dynamic, permitting their ideal point to change from one term to the next. As shown in the appendix, we find that our approach performs as well as or better than alternatives and does a better job at recovering the second dimension.

Recorded council votes were retrieved from city government websites. Chicago council votes are embedded in minute books. Page images of the minutes of meetings held since July 1981 have been scanned to PDF format and posted on the city's website. We used optical character recognition (OCR) software to convert them to text, after which we extracted the roll-call votes, which are displayed in a standard format. We validated the digitized votes by comparing them with legislative reports prepared by University of Illinois at Chicago professor Dick Simpson since the May 2003 general election (Simpson n.d.). This revealed 374 correctly captured voting events and nine missing voting events (2.3 per cent), which were added to the dataset, during this period. The complete 1981–2019 dataset contains 2,450 recorded votes.

Records of Toronto council votes starting with the January 2009 meeting are available from the city's open-data portal. Prior to 2009, individual councilors' votes were recorded in council minutes posted as PDF documents. We scraped these from the PDFs and integrated them with the post-2009 table. The resulting dataset contains 16,542 votes from the amalgamated city's first council meeting on January 2, 1998, to the most recently available meeting of the 2018–22 term (February 2, 2022).[2]

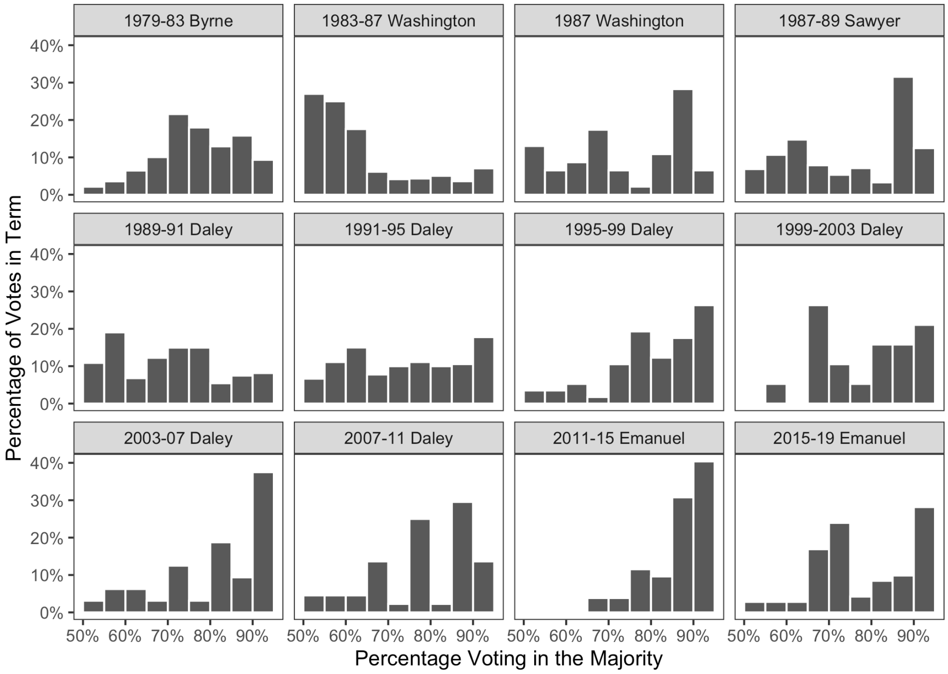
Figure 1 displays histograms of the distribution of vote margins in each council term, excluding those (nearly) unanimous votes in which more than 95 per cent of non-absent councilors voted with the majority. The right-most bar in each subgraph indicates the proportion of recorded votes that were almost unanimous; the left-most bar indicates the proportion that was nearly tied.[3] In Chicago, the 1987–91 term is divided into three segments defined by the mayoralty. Harold Washington died in office soon after his 1987 reelection and alderman Eugene Sawyer was appointed to take his place. He was defeated in a 1989 special election by Richard M. Daley, who retained power in the 1991 general election. The figures suggest that contestation was greater in earlier years in both cities; the proportion of lopsided votes within each term has increased over time.

[2] The datasets contain all votes on agenda items. It might seem reasonable to use only final votes, however we found that 10–15% of the time, the final vote is not the most contentious one in the sequence of votes on a particular proposal in Toronto. It was not feasible to distinguish between final and other votes in Chicago without manually examining the vote events in the minute books, which comprise over 875,000 pages.

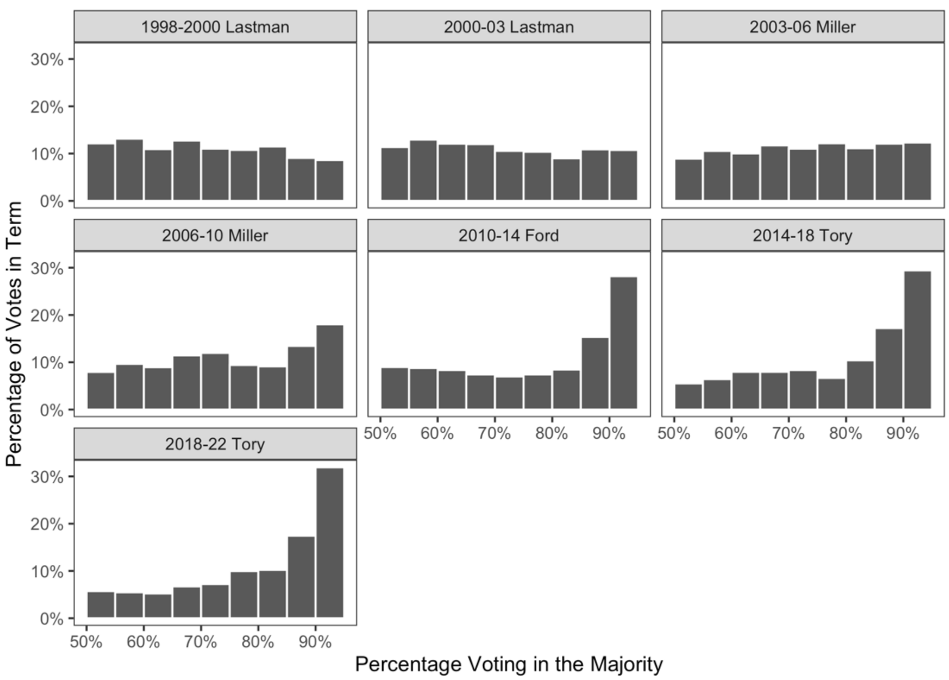
[3] Removing the near-unanimous votes compensates for inconsistency over time in the use of voice votes, which are not included in the roll-call data, to unanimously approve items. In Toronto, Mayor Rob Ford (2010–14) insisted that all decisions be recorded. While in later terms voice votes are once again used for many consensus decisions, many decisions that would have been made by voice vote in earlier terms are now recorded, inflating the number of unanimous votes (Consultation with City of Toronto's Clerk's Office, 1 Dec. 2020). We were unable to reliably extract voice votes from the meeting minutes.

Figure 1: Distribution of recorded council votes by degree of unanimity

Chicago



Toronto



Note: The histograms exclude vote events in which more than 95 per cent of councilors present voted with the majority.

<B>Ward Characteristics

To identify the meaning of the dimensions of disagreement, we correlated them with four categories of ward characteristics: the relative presence of *groups* defined by ethnic origin, race, and religion; markers of *status*, including income, educational attainment, occupation, and housing tenure and cost; *place qualities*, indicated by housing type and age, mobility, amenity accessibility, population density, and distance from City Hall; and *lifestyle*, including marital status and the presence of children, youth, and seniors. The data and their sources are described in the supplementary appendix. A correlation rather than a regression approach is used because wards are understood as bundles of partially correlated attributes – for example, racialized poverty – which, taken together, define the poles of each dimension.

<A>Findings

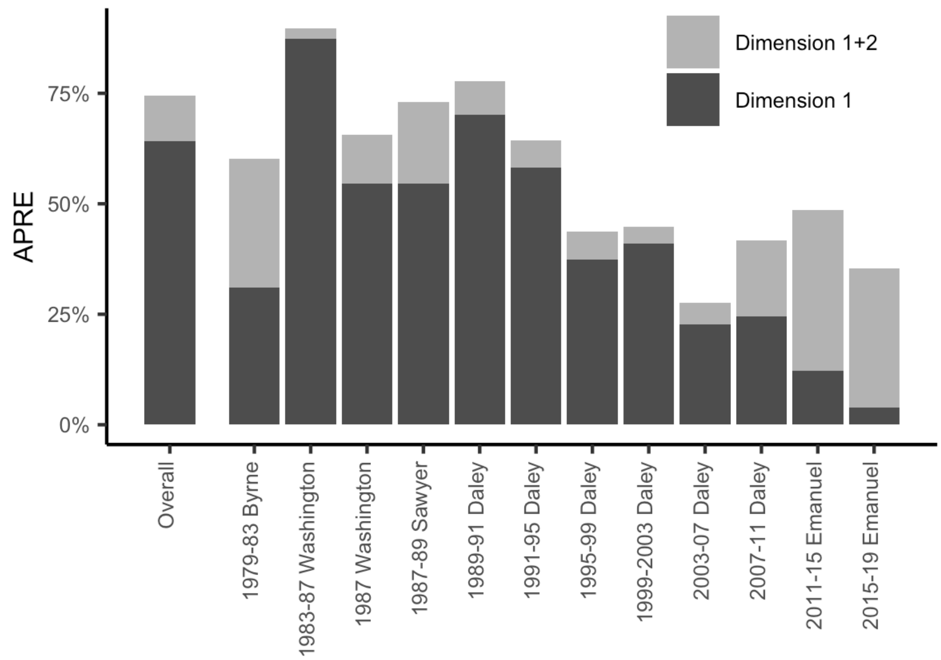
The analysis proceeds in three steps. First, following Poole and Rosenthal (2007, 36–37), we present aggregate proportional reduction in error (APRE) values associated with each dimension, which indicate the degree to which the dimension predicts councilors' votes. Second, we investigate the stability of councilors' positions over time as an indicator of enduring coordination on the councils. Finally, we identify the meaning of the dimensions by correlating them with ward characteristics, and, in Chicago, with councilor race.

<B>Low dimensionality

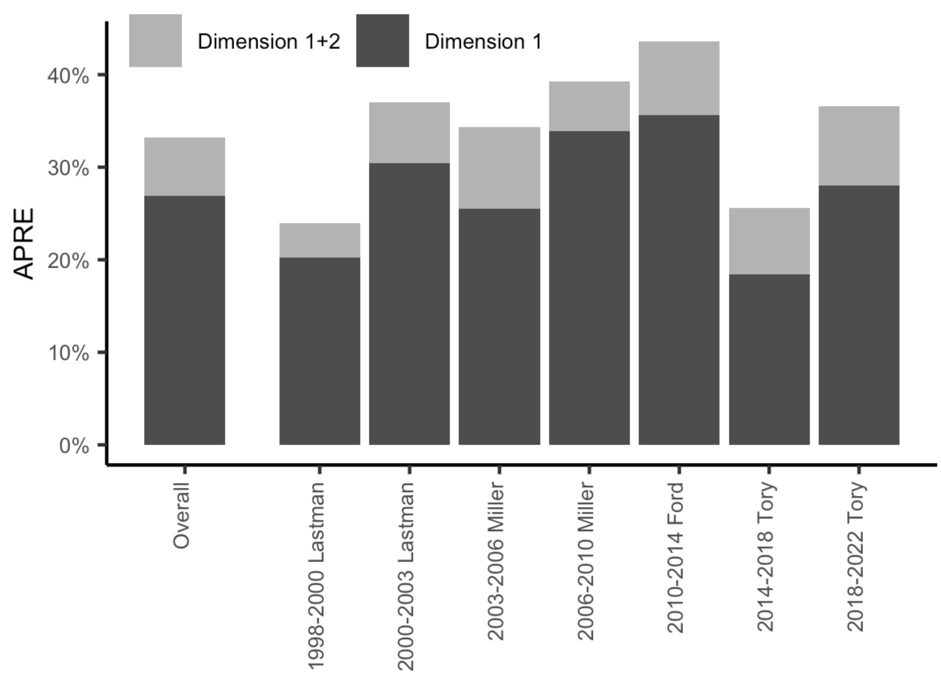
Our first finding is that councilor voting is low-dimensional in both cities. **Figure 2** presents the APRE of each dimension. A null model (i.e., one without any independent variables) can correctly classify the proportion in the modal category. The APRE gives the proportion by which classification errors are reduced with a model including the first (and second) dimension scores relative to the null model. The APRE is bounded on the high end at 1 but has no constant lower bound; PREs can be negative when the model predicts fewer cases correctly than the null model. Higher values and especially those close to one indicate that the votes more closely align with the dimension(s) in question.

Figure 2: APRE of two-dimensional models

Chicago



Toronto



In Chicago, there were 107,710 votes of which the null model correctly predicts roughly 80%. This is consistent with the considerable number of lopsided votes. The first dimension correctly predicts around 93% of the votes and the first and second dimensions together predict roughly 95% of the votes. This results in an APRE of 75%. Calculated term-by-term, the PRE diminishes over time. It peaks at 90% during Harold Washington's first term (1983–87), remains between the 65% and 76% through the end of Daley's first four-year term (1991–95). After 1995, it markedly diminishes, ranging from 28% in the 2003–07 term to 50% in 2011-2015. The relative importance of each dimension also varies over time. The first dimension does most of the work in most terms of office until the end of the 2003–2007 term; it accounts for between 75% and 97% of PRE. The 1983–87 term is virtually unidimensional. The second dimension becomes more important in later terms, and especially during Daley's final term (2007–11) and Emanuel's two terms (2011–19).

Toronto paints a similar picture, however the accuracy is lower. There are 598,658 votes to be predicted in Toronto, of which 493,199, or 82%, are correctly predicted by the null model. The first dimension correctly classifies 83% of the votes and the addition of the second dimension boosts that figure to 85%. This leads to a proportional reduction in error of around 33%. Term-by-term, error reduction is highest in Ford's term (2010–14), at 44%, and lowest in the first term of the newly amalgamated city (24%). The second dimension is most predictive, relative to the first, in Tory's two terms (2014–present).

The time-pooled two-dimension APRE values fall within the same range as found by Shor and McCarty (2011, 534) their analysis of U.S. state legislatures – between 28.3 and 82.7% – with Chicago's council (75%) scoring similar to Michigan (74%) and Toronto (33%) scoring similar to Nebraska's nonpartisan legislature (35%).

We draw three main conclusions from the classification analysis. First, roll-call voting in the two cities is low-dimensional. This suggests that there are enduring opposing groups of councilors, and that these groups endure despite the entry and exit of individuals from the council. Second, low-dimensionality is not unidimensionality. The second dimension is important in both cities, and matters a great deal within specific terms of office, sometimes eclipsing the first dimension. Finally, variation over time in the proportion of votes correctly classified reminds us that these are relatively unstructured legislative environments. There is a lot of "noise" in the voting, and more noise in some terms than in others. We further assessed this by examining the agreement scores – the percentage of times each pair of councilors agrees, call this a_{ij}, where i and j index two different councilors. The more structured the voting, the more we should see agreement scores at the poles (i.e., zero or one). We transform a_{ij} such that $a_{ij} = 1 - a_{ij}$ if $a_{ij} < .5$. This means that a_{ij} will range from 0.5 to 1 where numbers closer to one are consistent with structured voting and numbers close to 0.5 with unstructured voting. We found that the median agreement score was 84% in Chicago and 73% in Toronto. Toronto's lower score indicates that its council is slightly less structured than Chicago's. This is unsurprising given that Chicago's elections were partisan prior to 1999, and that Chicago's mayor has stronger prerogatives than Toronto's.

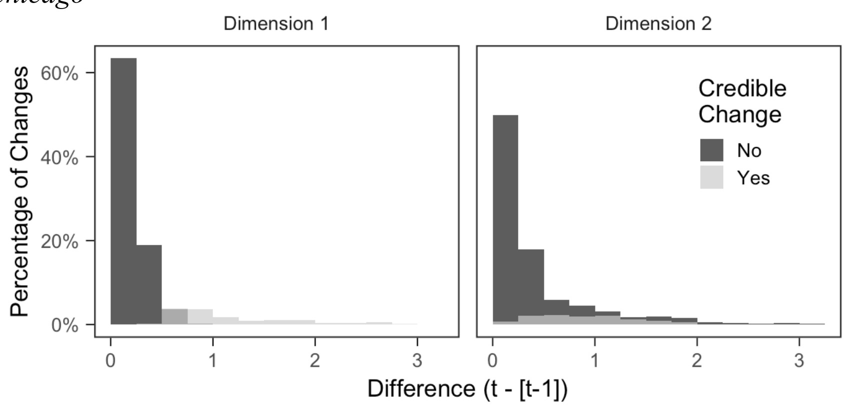
<B>Further evidence of durable coordination

A key advantage of our technique is that it enables councilors' positions on the dimensions to change from one term of office to the next. If most councilors' positions are stable over time, this is further evidence that coalitions among councilors are not transitory and issue-based, as predicted by pluralism, but enduring. **Figure 3** shows histograms of differences in councilor placement from one term of office to another for councilors who served in at least three terms.

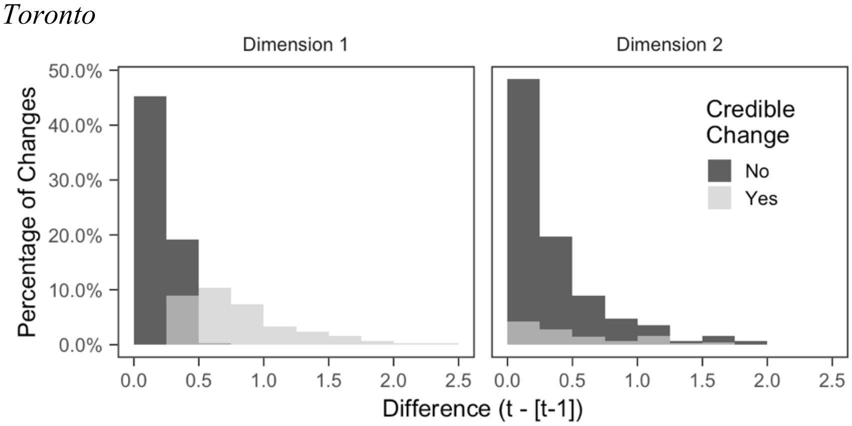
(The results differ little if we include councilors who served in at least two terms.) Because the underlying model is Bayesian, there is no single estimate of this difference, however we obtain the posterior distribution of the change in position from each term to every other term in which a councilor served. For each of these differences we calculate the posterior probability of a change and mark it as "credible" if the posterior probability of a change was greater than 0.9.

Figure 3: Mean inter-term movement in ideal points

Chicago



Toronto



Note: The horizontal axis indicates the mean inter-term movement in position on the dimension between terms, expressed in dimension units. The vertical axis indicates the proportion of councilors in each bin. Dark bars indicate movement smaller than a credibility test; light bars indicate that the movement is credible.

Applying this test, we find that while two-thirds of councilor movements met the credibility threshold on the first dimension in Toronto, one-third or less did on Toronto's second dimension and on both dimensions in Chicago. Only a small proportion of councilors moved by more than 0.5 units on any dimension. Some of this movement may be explained by changes in mayor. Alignment of priorities may pull councilors toward the mayor, while disagreement may push them away. Similarly, the arrival of sympathetic or antagonistic colleagues through electoral turnover may cause councilors' ideal points to shift over time. Nevertheless, the degree of stability observed suggests that councilors are anchored by strong political and policy predispositions, responsiveness to the distinct interests of their wards, or both.

<B>The meaning of the dimensions

Our final step is to assess what motivates the formation of durable alignments of legislators in the two cities. Given that ward councilors must please their constituents to be re-elected, we may expect the dimensions of disagreement to be associated with the characteristics of their wards, which in turn generate distinct sets of policy demands. To test this, we correlated ward characteristics with the two dimensions, both for each term separately and pooled across all terms. **Figures 4** and **5** show the correlation coefficients for Chicago for selected ward characteristic variables; **Figures 7** and **8** do the same for Toronto. Coefficients are displayed if they are statistically significant ($p < 0.05$), have absolute magnitude of at least 0.2, and appear in at least five terms of office in Chicago and four of the six terms in Toronto for Dimension 1 and two of the six terms for Toronto's much less temporally coherent second dimension.

Chicago. The correlation coefficients indicate that both of Chicago's dimensions are anchored by antagonistic group identities – Black vs. non-Hispanic white – however differences between the two dimensions in the other correlates suggest that race is inflected by socio-economic status and aspects of place and location within the city. Wards disproportionately inhabited by white "ethnics," immigrants, and people with classically "suburban" lifestyles – higher levels of home ownership, car-commuting, and traditional marriage – are positively associated with the first dimension. Wards with concentrations of African Americans, transit users, and young adults are associated with the opposite pole of the first dimension, as is the prevalence of apartment housing and markers of disadvantage: higher rates of divorce, unemployment, and vacant housing.

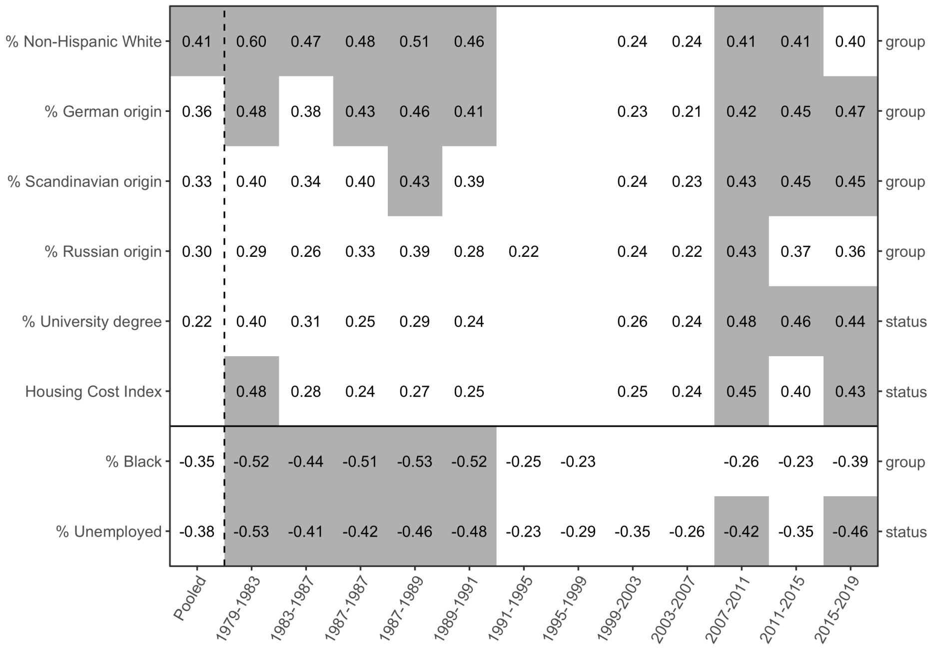
Figure 4: Chicago – Correlation of selected ward characteristics with dimension 1

	Pooled	1979-1983	1983-1987	1987-1987	1987-1989	1989-1991	1991-1995	1995-1999	1999-2003	2003-2007	2007-2011	2011-2015	
% married	0.57	0.48	0.67	0.75	0.70	0.71	0.69	0.70	0.63	0.59	0.55		lifestyle
% Polish origin	0.53	0.38	0.65	0.73	0.71	0.66	0.65	0.59	0.48	0.45	0.39		group
% Non-Hispanic White	0.52	0.25	0.68	0.77	0.71	0.68	0.69	0.70	0.49	0.50	0.25		group
% own home	0.45	0.27	0.48	0.68	0.65	0.59	0.59	0.45	0.59	0.61	0.27		status
% Italian origin	0.45	0.33	0.55	0.70	0.65	0.62	0.59	0.60	0.45	0.44	0.26		group
% Irish origin	0.39		0.54	0.65	0.58	0.51	0.58	0.59	0.40	0.40	0.23		group
% car commute	0.38	0.47	0.36	0.49	0.47	0.45	0.42	0.40	0.43	0.44	0.45		place
% immigrants	0.34	0.30	0.45	0.42	0.35	0.46	0.43	0.48	0.35	0.25	0.35		group
% Age 18-25	-0.27		-0.41	-0.48	-0.45	-0.30	-0.28	-0.22	-0.26	-0.38	-0.48	-0.33	lifestyle
Transitscore	-0.29	-0.36	-0.29	-0.37	-0.35	-0.30	-0.27	-0.22	-0.30	-0.34	-0.43	-0.29	place
% Apartments	-0.29	-0.24		-0.38	-0.36	-0.31	-0.27		-0.38	-0.44	-0.46	-0.22	place
% divorced	-0.39	-0.45	-0.58	-0.54	-0.45	-0.60	-0.59	-0.62	-0.41	-0.33			lifestyle
% Dwellings vacant	-0.41	-0.21	-0.23	-0.56	-0.51	-0.48	-0.49	-0.43	-0.65	-0.65	-0.48	-0.21	place
% transit commute	-0.45	-0.58	-0.53	-0.61	-0.60	-0.62	-0.59	-0.56	-0.42	-0.45	-0.39		place
% Black	-0.54	-0.35	-0.70	-0.72	-0.64	-0.73	-0.73	-0.74	-0.52	-0.43	-0.36		group

Note: Pearson correlation coefficients are significant to a level of $p < 0.05$. Coefficients smaller than ±0.2 are not shown. Coefficients larger than ±0.4 are shaded to indicate greater magnitude. Coefficients are ordered from largest to smallest on the "Pooled" column.

The second dimension is also anchored by race and ethnicity, but inflected by socio-economic status rather than by place characteristics. Wards positively associated with the second dimension are disproportionately home to non-Hispanic whites, those with high educational attainment, and those who inhabit high-cost housing. Wards negatively associated with the second dimension are disproportionately inhabited by African Americans and the unemployed.

Figure 5: Chicago – Correlation of selected ward characteristics with dimension 2



Note: Pearson correlation coefficients are significant to a level of $p < 0.05$. Coefficients smaller than ±0.2 are not shown. Coefficients larger than ±0.4 are shaded to indicate greater magnitude. Coefficients are ordered from largest to smallest on the "Pooled" column.

In addition, we examined the relationship between councilor race, ward characteristics, and councilor voting. Councilor race and their wards' racial composition are strongly correlated. Across all years in wards represented by African American councilors, the mean percentage of African Americans is 87%, Latinx 7%, and whites 5%. In wards represented by white councilors, the mean percentage of whites is 59%, Latinx 21%, and African Americans 13%. While the number of African American aldermen on Chicago's 50-ward council has ranged between 16 and 19 over the study period, the number of white alderman has steadily decreased from 34 to 21 as growing numbers of Latinx and Asian candidates have been elected. **Figure 4** displays boxplots of the dimension scores for African American and white aldermen. The figure indicates that an alderman's position on Dimension 1 is strongly associated with her race. During the mayoralties of Harold Washington (the city's first Black mayor), Eugene Sawyer, and Richard

M. Daley, which bracket a period of racially polarized conflict known as the “council wars” (1983–86) and its aftermath, the interquartile ranges of African American and white aldermen’s positions barely overlap. Starting in the early 2000s, African American aldermen move closer to white aldermen. During Emanuel’s two terms, white aldermen converge with their African American counterparts. The relationship been alderman race and the second dimension is less pronounced. African American and white aldermen’s positions on the second dimension significantly diverge most strongly following the “council wars” period during the Washington’s abortive second term, his appointed replacement Sawyer’s interregnum, and Daley’s first term following the 1989 special election.

Figure 6: Chicago – Dimension scores by race of councilor by term of office

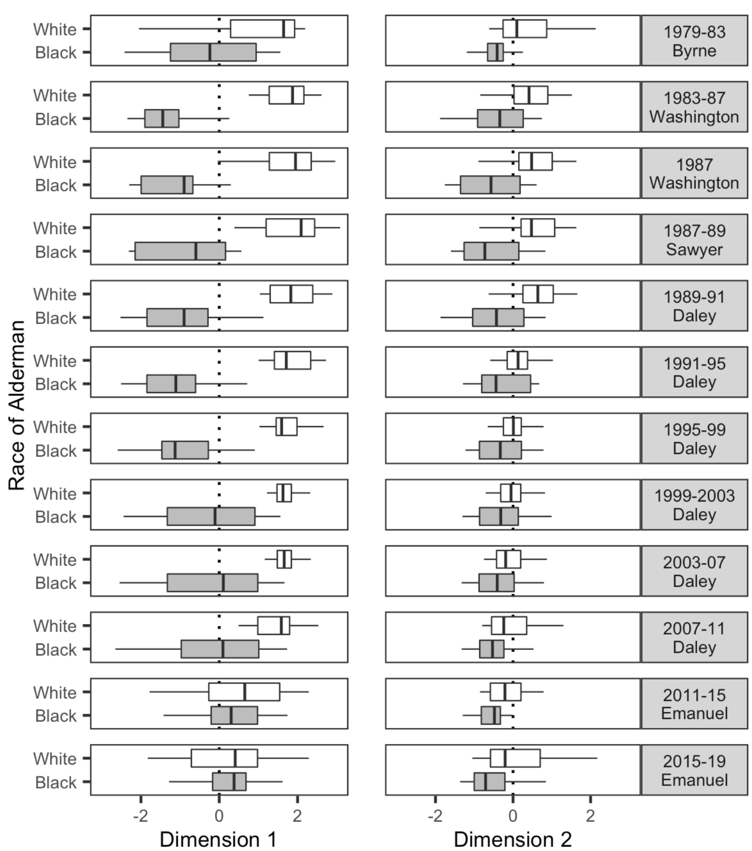
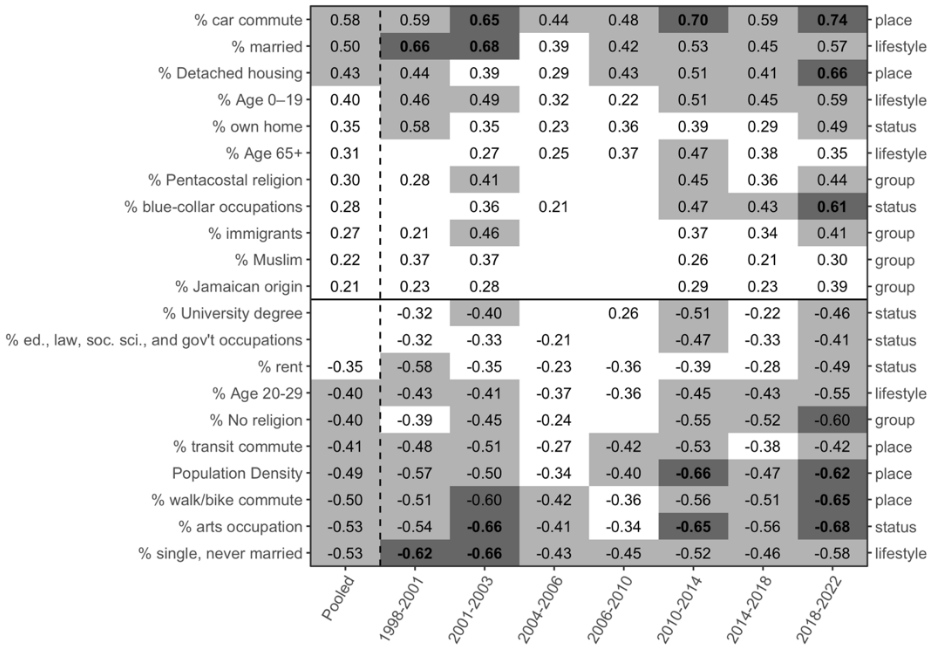
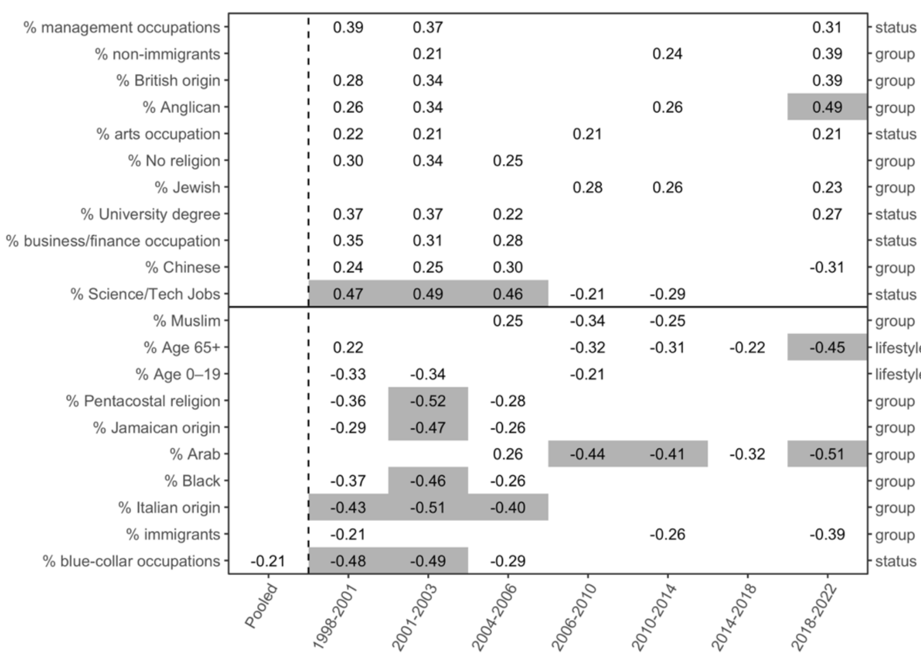


Figure 7: Toronto – Correlation of selected ward characteristics with dimension 1



Note: Pearson correlation coefficients are significant to a level of $p < 0.05$. Coefficients smaller than ±0.2 are not shown. Coefficients larger than ±0.4 are shaded to indicate greater magnitude. Coefficients are ordered from largest to smallest on the "Pooled" column.

Figure 8: Toronto – Correlation of selected ward characteristics with dimension 2



Note: Pearson correlation coefficients are significant to a level of $p < 0.05$. Coefficients smaller than ±0.2 are not shown. Coefficients larger than ±0.4 are shaded to indicate greater magnitude. Coefficients are ordered from largest to smallest on the "Pooled" column.

It appears, then, that aldermanic voting remained polarized by the racial makeup of wards during the 2010s but not the race of the aldermen themselves. However, the growing importance of the second dimension, and the diminishing importance of the first, suggests that the racial divide is in recent decades inflected by the socio-economic status of ward residents as opposed to place characteristics.

Toronto. Turning to Toronto, **Figures 7** and **8** suggest that the first dimension is primarily associated with place-based characteristics and associated lifestyles. Centrally located, dense, and walkable wards with good transit access are positively associated with the first dimension. The inhabitants of these wards are disproportionately single, young, work in arts, entertainment, and recreation or management occupations, and commute on foot or by bicycle. Wards negatively associated with the first dimension tend to be located in the city's suburban periphery. Residents of these wards disproportionately commute by car and live in low-density, single-detached housing. The prevalence of married families and children are evidence of traditional domestic living arrangements. These wards also contain concentrations of immigrants and blue-collar workers. It seems that the first and dominant dimension of councilor voting behavior is related to a distinction between two spatially distinct urban conditions: a postindustrial, "youthified," densifying, walkable, gentrifying core versus a car-dependent, low-density residential periphery that is oriented toward domestic lifestyles.

The second, weaker dimension may represent something quite different: a status distinction rooted in ethnicity and occupation. Wards positively associated with the second dimension are disproportionately home to "ethnic" communities and blue-collar workers. Positively correlated wards are associated in some years with the prevalence of British-origin residents, "high church" Anglicans, and atheism, and also high-skilled professional workers.

Taken together, this may reflect a status distinction between Toronto's established "anglo" economic and cultural elites and those at the bottom of the economic ladder, many of whom are immigrants or the children of immigrants.

Discussion. The correlation analysis reveals that cooperation and conflict among elected elite is social-spatial in both cities; that is, the dimensions' poles correspond to recognizable bundles of ward characteristics. The dominant legislative cleavage in each city is not ideology. Chicago's first dimension is the nexus of *race and place*, whereas the second dimension is the nexus of *race and status*. That race anchors both dimensions despite the dimensions being orthogonal to one another is a significant finding, one that reinforces Hajnal and Trounstine's (2014) conclusion that racial conflict is foundational to the politics of American central cities. The analysis of Chicago's dimension scores by race of alderman adds additional evidence that the voting behavior of municipal legislators has been durably divided on the basis of race, although this has diminished over time. Explaining this attenuation would require qualitative research that is beyond the scope of this paper, however we hypothesize that the decline of the disciplining role of the Democratic Party "machine" (McClelland 2019), coupled with the declining demographic dominance of African-Americans and whites in favour of Latinx and residents of other ethno-racial backgrounds, has diminished a previously important structuring element. The emergence in the 2010s of formal council caucuses with overlapping memberships – Black and Latino Caucuses, as well as the Progressive Reform, Democratic Socialist, and LGBT caucuses – is symptomatic of the machine's decline, and each has provided a focal point for opposition to mayoral agendas (Simpson 2020).

The Toronto correlation analysis reveals a very different primary cleavage. The first dimension represents a distinction rooted in lifestyles associated with *core* and *periphery* locations within the city, each with distinct place qualities, material interests, and identities: the dense, gentrifying zone of young, single professionals and arts workers at the city's centre, juxtaposed with peripherally located wards characterized by automobile commuting, low-density postwar suburban built environments, and domestic lifestyles. The second dimension represents a status cleavage rooted in group differences. Wards at the positive pole are disproportionately home to those with higher status occupations and higher education. At the negative pole are low-status wards whose residents are disproportionately blue collar and belong to immigrant-basec communities. While the first dimension contrasts different locations within the city, the second contrasts established in-groups from out-groups.

<A>Conclusion

The goal of our analysis is to identify the underpinnings of legislator behavior in two relatively unstructured institutional environments: nonpartisan city councils in Chicago and Toronto. To do so, we employ an innovative combination of methods to analyze recorded votes on legislative proposals in a way that prejudges neither their dimensionality nor the dimensions' meanings. Our approach positions councilors on a common scale for each dimension across the multiple terms of office spanning multiple decades while at the same time permitting their positions to change over time. We then investigate the meaning of the dimensions, and therefore the political divisions that structure local politics, by correlating them with ward characteristics and, in Chicago, the race of aldermen.

The analysis yields two major findings. First, we find that disagreement on Chicago's and Toronto's councils is neither high-dimensional nor no-dimensional, as might be predicted by a pluralist analysis. We find that two latent dimensions predict a substantial proportion of councilors' votes on legislative proposals in both cities. Coordination among councilors is persistent, as evidenced by the spread between and minimal change in their ideal points over time. This suggests that councilors have durable orientations toward these dimensions. These divisions persist over multiple terms of office spanning decades, transcending the careers of individual politicians. This finding is an important contribution to the study urban politics, as it indicates that durable political divisions exist among municipal elected elites even in the absence of parties and other disciplining institutions.

Our finding that disagreement is low-dimensional does not lead us to fully dismiss the structuralist expectation of convergent preferences and consensual politics at the elite level. After all, we find that many items are passed or defeated by lopsided margins and unanimous voice votes (which could not be incorporated into our database) are commonplace. Without knowing the topic of the motions being voted on, we cannot determine which issues lend themselves to division and which to convergence. Adopting Peterson's (1981) typology of policy arenas, it may be that councilors converge on developmental issues, as predicted by the structuralist perspective, while potentially diverging on redistributive and allocational policy questions. For example, we might expect councillors to unanimously support the setting of the property tax rate, but find that "core" councilors may support, and "periphery" oppose, spending on bicycle infrastructure or public art projects. Incorporating issue types would be a potentially fruitful, if highly labor-intensive, extension to the analysis.

Our second major finding is that councilor voting in both cities has socio-spatial foundations; it is systematically aligned with wards characteristics. This is evidence that elected representatives are responsive to the group interests and identities present in their wards. However, the nature of the groups differs in the two cities. In their comprehensive analysis of exit polls and election results in American city elections, Hajnal and Trounstine (2014) find that race is the dominant determinant of vote choice:

> The urban electorate is shaped in part by class, religion, sexuality, age, gender, and a host of other demographic measures, but race seems to be more central and more decisive than all of these other factors. … Most accounts of politics at the local or national level point to party identification or ideology as the main driving forces in American politics …. But the results presented here suggest otherwise. … [In] local democracy, it is race more than anything else that tends to dominate voter decision making. (86)

This claim is supported in the Chicago case, where we find that race and ethnicity manifest in orthogonal dimensions otherwise defined by place characteristics and socio-economic status.

We find a very different pattern in Toronto despite considerable ethnic, racial, and religious diversity. Legislative disagreement divides primarily on legislators' wards' built environment characteristics and associated lifestyles and economic activities, pitting representatives of the high-density gentrified core against those of the car-dependent residential periphery. A secondary, status-based division between representatives of wards with different occupational, educational, and ethnic profiles is weaker. This is consistent with the findings of Doering, Silver, and Taylor (2021), who find similar relationships between neighborhood-scale election results and the social and physical characteristics of neighbourhoods in the same cities.

Our finding that legislative disagreement is multidimensional and that the meaning of these dimensions differs in two large North American cities has important implications for the study of urban politics. At the national scale, Norris and Inglehart (2019) argue that postmaterial value change has disrupted the earlier grounding of political competition in "polarization over welfare distribution and the state role in the economy" (50) – the left/right divide. In its place, they propose a two-dimensional model featuring two orthogonal cultural cleavages, the first dividing pluralists from populists, the second dividing authoritarians from libertarians. We may also expect multidimensionality for the same reason in local politics, but tailored to matters distinct to urban life and urban conditions, such as commuting, the quality of the built and natural environment, and neighborhood change. While not focused on electoral politics and representation, Urry (2004) identifies automobility with individualism and privatism, which joins with the work of Castells (1977) and Fischel (2005) to suggest that, beyond race and class, political interests, and perhaps enduring political cleavages, may emerge from patterns of consumption. This is consistent with theoretical and empirical inquiry into the development of postmaterial urban social movements (Castells 1983; Offe 1987), and it is therefore not surprising to see durable cleavages within cities that are rooted in place-based lifestyles associated with distinct built environments and the material interests and identities they generate. Future research could build on this observation by theorizing and typologizing such group identities and interests and examine their activation in electoral campaigning.

An additional extension would be to investigate mechanisms of representation – what connects ward characteristics to decision making by their representatives. It may be that councilors are sensitive to their wards' distinct social composition and material interests. Alternatively, councilors may be products of their ward's social and political milieux, and therefore bring associated attitudes into the legislative body. Surveys of or interviews with present and former councilors could unpack the causal connection, as could ward-level surveys comparing the policy preferences of elected officials and residents.

Much can be gained by applying our methodological approach to additional cities for which the necessary data are available. We may find that the relationship we found in Chicago and Toronto between the dimensionality of legislator voting and the characteristics of the areas they represent breaks down when wards are very large, elections are at-large rather than by ward, or in cases where strong discipline is imposed by political parties or by mayoral institutional prerogatives. Ultimately, we argue that systematic attention to the decisions municipal councilors make has much to offer urban political studies, especially as the construction and analysis of large-scale roll-call databases has become technically feasible.

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SUPPLEMENTARY APPENDIX

Political Divisions in Large Cities: The Socio-Spatial Basis of Legislative Behavior in Chicago and Toronto

Appendix A: Technique for the dimensional reduction of legislative roll-call votes

We rely on several methodological tools to estimate the dimensionality of legislative voting. The main model of interest is the Bayesian dynamic item response model popularized by Martin and Quinn (2002) in their analysis of US Supreme Court justice ideology over time. Unlike them, we are interested in estimating models in potentially several dimensions. This poses several challenges. First, classical Bayesian modeling using BUGS, JAGS or Stan does not scale well to the number of votes in Chicago or Toronto. Thus, using these tools would be quite slow. Imai, Lo, and Olmsted (2016); (Imai, Lo, and Olmsted 2020) have developed a scalable algorithm to produce Bayesian IRT results. The challenge here is that these routines have only been coded for models with a single, underlying latent dimension. Both DW-NOMINATE (Poole 2005) and Potthoff (2018) principal components analysis (PCA) solution allow evolution in ideal points over time, but they are constrained to a pre-determined parametric functional form. We have no reason to believe that the evolution of councilor ideal points will even be monotonic, much less linear over time and thus these solutions are not sufficiently flexible for our purposes.

While Pothoff's solution is not appropriate for our purposes, the general idea behind it led us in the direction of a workable model. We adopt a multi-stage strategy that allows us to use existing unidimensional Bayesian D-IRT models to estimate multiple dimensions. The general idea that we take from Potthoff is to use a PCA (or a similar tool) on the councilor votes roll-call matrix, R. However, we do this in an effort to identify votes that are likely to be associated with one of potentially several underlying sources of variation. For any particular vote y_j, we want to know whether it aligns with dimension $1, 2, \dots, m$. We will then use this information to estimate m different uni-dimensional D-IRT models that, taken together, will provide multi-dimensional estimates of councilor ideal points in space and time.

We first detail the steps taken in the analysis, which are implemented in an R package called [package name excised for peer review], available from [URL excised for peer review]. We then show how the results we generate stack up against several different NOMINATE-based solutions.

Step 1: Preliminary estimation of latent variables. Instead of using PCA, we use a generalized low rank model (GLRM), which is more permissive of missing data than PCA (Boehmke and Greenwell 2020; Udell et al. 2016).[1] Our roll call matrix R has two types of missing data present: *structural* missing data that occur when a councilor did not serve in a particular term, and *idiosyncratic* missing data that occur when serving councilors are absent from voting on particular items. In Toronto, of the 2,067,750 cells in the matrix, 598,658 votes are observed. In Chicago, 107,710 out of 477,046 votes are observed.[2]

PCA uses a singular value decomposition to decompose an observed rectangular data matrix in the following way:

$$R = UDV'$$

Where U is an ortho-normal matrix of left singular vectors that provides information about the independent sources of variation in R. The V matrix comprises the right singular vectors and provides information about the rows that relates to the independent sources of variation in R. Finally, D is a diagonal matrix giving the amount of total variance captured by each of the underlying dimensions. The V matrix is the matrix of coefficients that maps the observed data R onto the principal components; it is the matrix that creates an orthogonal, variance maximizing, linear combination of the R matrix. This model assumes linear relationships between the observed data and the components.

By contrast, the GLRM does not assume any particular functional form of the relationship. Rather than relying on singular value decomposition, it creates the following decomposition

$$R = XY$$

where X is an $n \times m$ matrix that identifies the value for each of the n observations on each of the m *archetypes*, which are analogous to the components from PCA. Y is an $m \times k$ matrix that gives the values of each of the k observed variables on each of the m archetypes. In this model, categorical features are approximated by minimizing the multinomial loss function. First, categorical data are "one-hot encoded" where all levels of the category are made into effect coded variables given a value of 1 for those observations that are in the category represented by

[1] Potthoff (2018) proposes an imputation strategy that generates a single imputation based on existing data. We are uncomfortable imputing 30% of the data to generate the ideal points and instead use a different strategy.

[2] In Toronto, 151,647 votes (7% of all votes, 10% of missed votes) were missed due to absence; the average councilor misses roughly 25% of votes. The remaining missed votes come from councilors who do not serve in all council terms. In Chicago, 22,238 votes (5% of all votes, 6% of missed votes) were missed due to absence, with the average councilor missing around 15% of votes.

the variable and –1 for all others. Then, through alternating minimization (of X and Y), the following loss function is minimized:

$$L(u_{ij}, a_{ij}) = \log\left(1 + \exp(-a_{ij}u_{ij})\right)$$

where a_{ij} are the effect coded variables and u is $x_i y_j$, the inner product of the appropriate row and column archetypes.

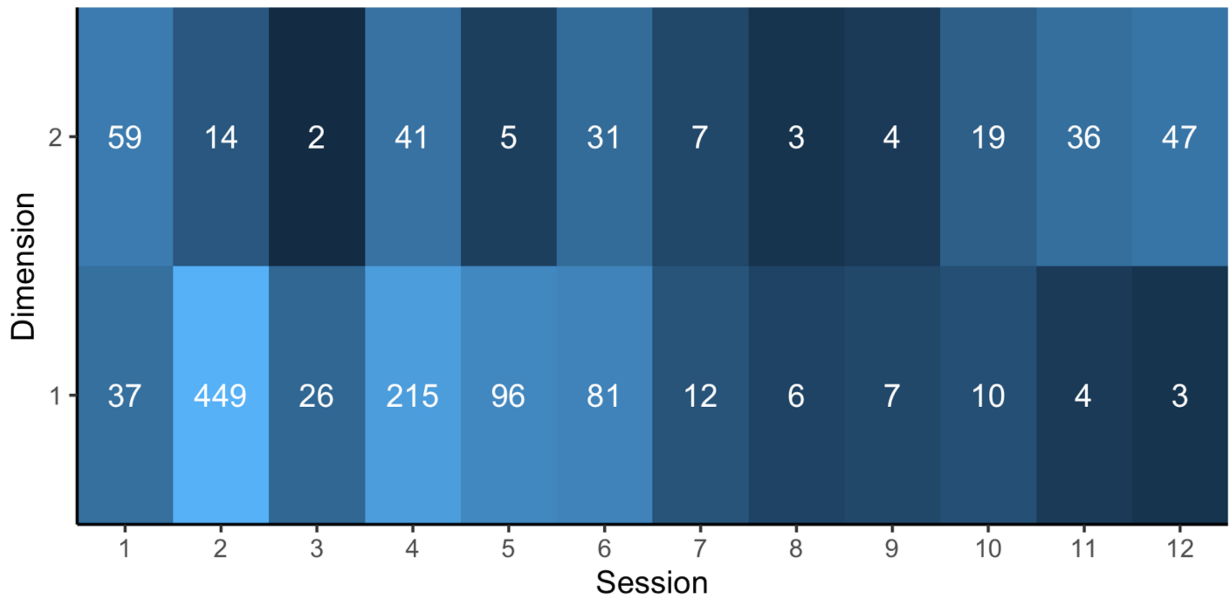
Ultimately, the X matrix from the GLRM serves the same purpose as UD, or equivalently RV, from the PCA, however without requiring the imputation of missing data. If the underlying PCA model of linear relationships holds, then the GLRM produces the same result as PCA. However, when non-linearities are present, the results will not necessarily be ortho-normal as they are in PCA. As this is the case in which we find ourselves, we perform a PCA on the X matrix from the GLRM to produce a variance-maximizing orthogonal transformation of X, call this $X^{(o)}$, with columns $x_l^{(o)}$ for $l = 1, \dots, m$.

In both of our applications, there are not enough votes to specify more than two dimensions. Thus, we have the GLRM produce two archetypes and we then use the PCA to orthogonalize the result. Theoretically, one could estimate any number of archetypes and then use PCA to reduce the dimensionality, but without considerable non-linearities and non-additivities, there would be little value in extracting more than a few archetypes. We use the H2O platform with R as an interface to estimate the GLRM.

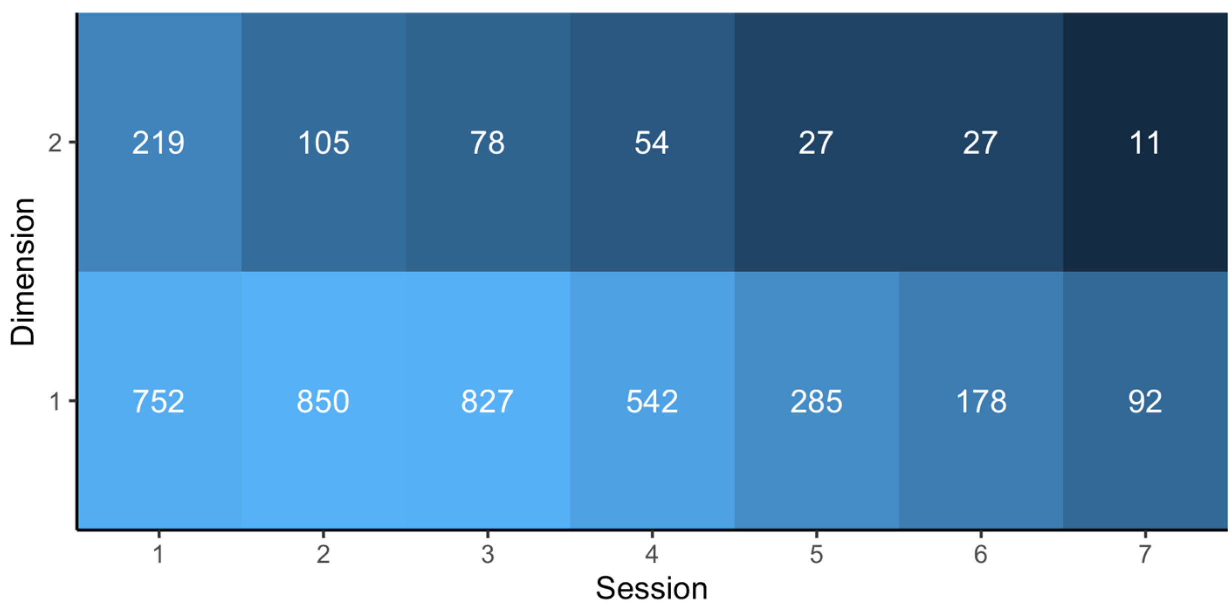
Step 2: Identify votes best predicted by the latent variables. Next, we subject the results of our GLRM to a PCA to order the columns from most to least variance explained and store the components in $X^{(o)}$. In our case, $X^{(o)}$ is an $n \times 2$ matrix. For each of the columns in $X^{(o)}$, we estimate a separate logistic regression model of r_j (roll call vote j) on $x_l^{(o)}$ (the l^{th} column from $X^{(o)}$). For each model, we calculate the PRE – the proportion by which prediction errors are reduced by using the independent variable to model the predicted probability of an affirmative vote. We use these PRE scores to identify the component that best predicts each variable. Specifically, we assign r_j to dimension l if $PRE_l > PRE_m \; \forall \; m \neq l$ and $PRE_m \leq 0.35 \; \forall \; m \neq l$. That is, we assign r_j to dimension l if the PRE for dimension l is bigger than the others and if none of the other PREs are higher than 0.35.

The value of 0.35 is arbitrary but chosen in our case for a particular reason. A value of 0 would be best, all else equal. That would suggest that r_j is explained to some degree by dimension l and not at all by any of the other dimensions. In our applications, this did not identify sufficiently many votes to estimate both dimensions. By increasing the cutoff, we acknowledge that votes will be a function of potentially both dimensions, though we assign the vote to the dimension that is a better predictor. To the extent this matters, it will tend to reduce the separability of the dimensions. That is, in a world where many of the votes are nearly equally well predicted by more than one dimension, those dimensions would be substantively quite similar. In our applications, this is not the case.

Figure A1: Number of votes assigned to each dimension in each term



a) Chicago



b) Toronto

The result of this step is to identify a set of votes that aligns with each dimension. **Figure A1** shows the number of votes assigned to each of the two dimensions in each term of office in the two cities. It is important to note that even though we use a relatively small number of votes to identify the dimensions we will ultimately evaluate the performance of the model's ability to explain *all* votes regardless of whether they were included in the model or not.

Step 4: Bayesian dynamic IRT for each dimension. We have now identified two sub-matrices of R: R_1, R_2 – each of which contains the subset of votes that is best predicted by each of the two dimensions identified above. We use each of these matrices in turn in a Bayesian D-IRT model.[3] The model is the same as used by Martin and Quinn (2002), where:

$$y_{ijt} = \Phi(\alpha_{jt} + \beta_{jt}\xi_{it})$$

Here, α_{jt} is like the difficulty parameter in a conventional IRT parameterization, but higher values of α indicate higher probabilities of voting in the affirmative, and thus "easier" items. β_{jt} is the discrimination parameter for vote j,[4] with prior $\begin{bmatrix}\alpha_{jt}\\ \beta_{jt}\end{bmatrix} \sim N\left(\begin{bmatrix}0\\0\end{bmatrix}, \begin{bmatrix}1 & 0\\ 0 & 1\end{bmatrix}\right)$. The councilor ideal points are assumed to follow a random walk:

$$\xi_{it}|\xi_{i,t-1} \sim N(\xi_{i,t-1}, \omega_i^2)$$

from the first to the last period that the councilor was in the data.[5]

[3] We estimate the Bayesian D-IRT model using the emIRT package in R (Imai, Lo, and Olmsted 2016, 2020).

[4] We mimic the notation in Imai, Lo, and Olmsted (2016), but note that the t subscript on the α and β parameters is irrelevant, because no vote occurs in more than one term.

[5] Several councilors had discontinuous tenures on Toronto's council. Palacio was appointed to fill Fontino's vacated seat in term 1. He did not run in term 2, but ran and was elected in terms 3 through 6. Both Saundercook (served terms 1, 3, and 4) and Moeser (served terms 1, 2, 4, 5 and 6) were defeated for one term in the middle of their tenures. For these councilors, we estimate the missing point in the middle of their tenure, but do not use it in our subsequent analyses.

Table A1: Prior identification of councilor ideal points

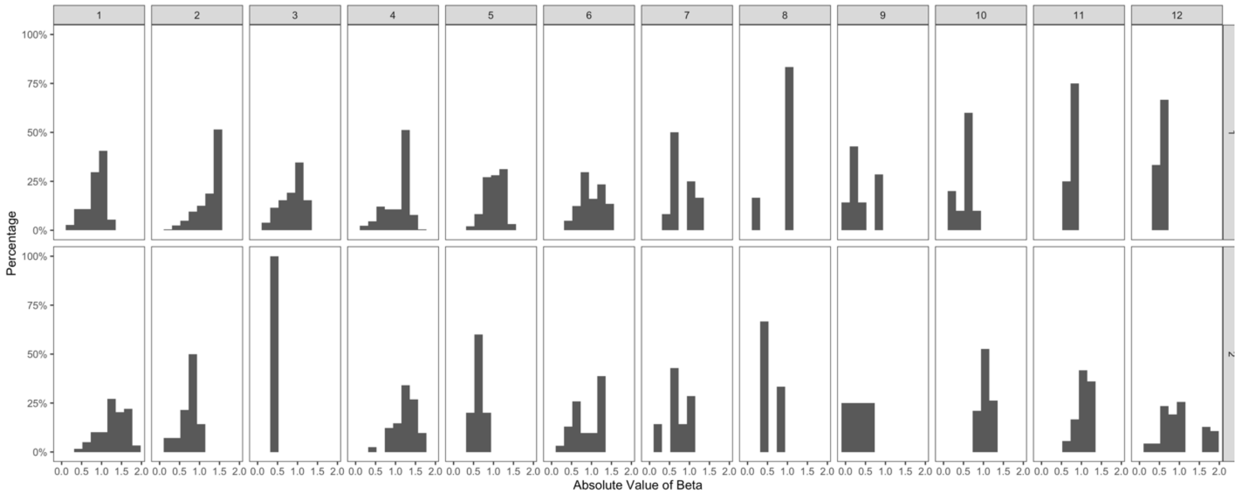
City	Dim	Negative	Positive
Toronto	1	Mihevc (served terms 1–6)	D. Holyday (served terms 1–5)
	2	Shiner (served terms 1–6)	Augimeri (served terms 1–6)
Chicago	1	Streeter (served terms 1-7)	T.W. Cullerton (served terms 1-6)
	2	Waguespack (served terms 10-12)	Brookins Jr. (served terms 9-12)

To identify the model, we must select councilors who inhabit positions toward the extremes of the dimension (though they need not be the most extreme councilors). We use an automated method to find the anchors. To do this, we find the two councilors who (a) voted on at least 50 proposals together and (b) agreed the least in their cast votes. We set these two councilors at opposing ends of the scale in each dimension and used our substantive knowledge of politics in each city to identify the appropriate pole in model post-processing. **Table A1** displays the identifying constraints for each dimension. For each of the councilors in the negative column we set their prior mean to –1 and the prior standard deviation to 0.1. For the councilors in the positive column, we set their prior means to 1 and the prior standard deviation to 0.1. We gave the remainder of the councilors standard normal priors. The prior standard deviation of the ω_i is set to 0.1 for each councilor. We supplied starting values for $\alpha_{jt} = 0, \beta_{jt} = 1$. For all councilors, we used starting values that were their values on the $X^{(o)}$ rescaled to the range (–2, 2) plus a draw from a random-uniform distribution in the range (–.25, .25).[6]

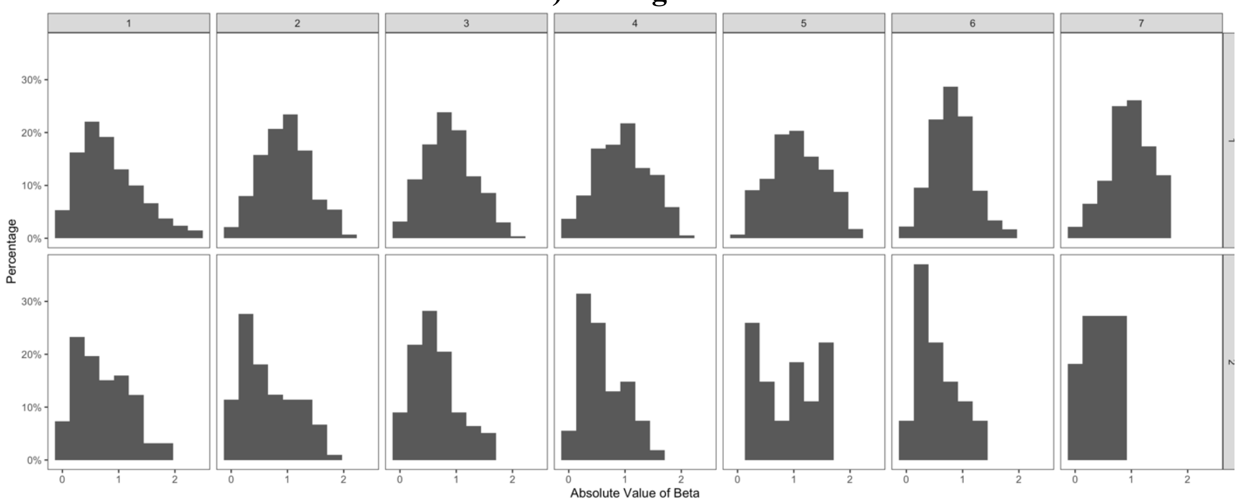
Diagnostics. **Figure A2** shows the absolute value of the β terms from the model. These describe the relationship between the latent variables and the observed votes. The first-dimension votes have most of their coefficients well away from zero, meaning that the observed votes reliably indicate the underlying dimension. While a greater proportion of the other dimensions' coefficients are close to zero in some terms, there is considerable variance in some terms, again suggesting that most votes are reliably associated with the underlying dimension.

[6] Recall that $X^{(o)}$ has only one observation per councilor.

Figure A2: Absolute values of β by term (column) and dimension (row)



a) Chicago



b) Toronto

Caption: The histograms above give the density of the absolute value of the discrimination parameter in the IRT. Higher values mean that the latent variable was more predictive of that particular vote. That is, the vote was a more reliable indicator of the underlying dimension.

Step 5: Orthogonalization. Since we estimate the latent dimensions independently, they are not uncorrelated as they might have been had we estimated (and imposed such a constraint on) a single multidimensional solution. To produce independent dimensions, we take the final step of orthogonalizing the latent dimensions. We do this by a Gram-Schmidt transformation. Thus, $\xi_1^{(o)}$, our estimate of the orthogonalized dimension 1 is simply ξ_1. We estimate the following equation to orthogonalize ξ_2:

$$\xi_2 = d_0 + d_1\xi_1 + \nu_1$$

The term ν_1 is the residual from this model that is uncorrelated with ξ_1 by construction; we set $\xi_2^{(o)} = \nu_1$. This will allow us to assess the independent relationships of these latent variables with various ward-level characteristics.

Step 6: Assign Polarity. Despite the priors used to identify the model, there remains the possibility that the model flips signs of the latent variables in any term. We start with terms 1 and 2. We calculate the Euclidian distance between the two vectors of values. We then flip term 2 (multiply it by –1) and calculate the Euclidian distance again. If the flipped distance is smaller than the original distance, we retain the flipped coordinates, otherwise we retain the original coordinates. We then move to terms 2 and 3 and repeat the process until we have worked through all sequential pairs of dimensions. There may remain some idiosyncratic polarity errors that can be diagnosed by visual inspection and then fixed by multiplying coordinates for that dimension in that term by –1.

To conclude, this procedure affords several advantages relative to other possible approaches. First, it locates legislators in multidimensional space on a common scale across time. Second, each dimension is orthogonal to the others, meaning that it potentially represents a distinct axis of disagreement. Third, each legislator's position is dynamic, permitting their ideal point to change from one term to the next in a very flexible way.

Comparison. We move to a comparison of our proposed method with three alternative approaches – DW-NOMINATE, one-shot W-NOMINATE where ideal-points for each councilor are fixed, and a term-by-term W-NOMINATE where the ideal points can change over time in a maximally flexible way, but the dimensions and positions are not necessarily comparable over time. We consider three sets of results: the squared correlation between our results and the NOMINATE-based results, the aggregate proportional reductions in error (APRE) for each model, and correlations of the results with contextual variables.

Squared correlation between results. In Toronto, the squared correlation between our first-dimension measure and the DW-NOMINATE measure is 0.83, and for the other W-NOMINATE measures (term-by-term and one-shot), the squared correlations are much smaller at roughly 0.14. This suggests that in some cases DW-NOMINATE may be a suitable alternative, but that the other two options are likely not appropriate. None of the second-dimension correlations are above even in magnitude – suggesting that each is picking up something quite different from the others. A similar pattern holds in Chicago, where our procedure's and DW-NOMINATE first dimension scores have a squared correlation that is lower, but still respectable at 0.57, but the correlations among the second-dimension scores are quite low – none rising above 0.03. The second-dimension correlation between our procedure and DW-NOMINATE scores is approximately zero.

Table A2: APRE Scores for Two-Dimensional models

	Toronto	Chicago
One-shot W-NOMINATE	0.28	0.13
Term-by-Term W-NOMINATE	0.35	0.76
DW-NOMINATE	0.35	0.75
Our procedure	0.32	0.75

APRE comparison. **Table A2** shows the APRE scores for the two-dimensional models in both Chicago and Toronto. In Toronto, the DW-NOMINATE model does slightly better than our solution. The one-shot W-NOMINATE model is not as good a predictor of votes, but the fact that it is not that much different than the dynamic models suggests that ideal points may not vary so much over time. Chicago tells a substantively similar, though more extreme story. Here, all three dynamic models produce roughly the same APRE, though the DW-NOMINATE and our procedure produce results that are comparable over time at the expense of very little extra predictive power. In general, in two dimensions, our model preforms similarly to the DW-NOMINATE model in terms of predictive power.

Correlation with substantive variables. Finally, we can compare the substantive interpretation offered by these other methods. Rather than comparing across all models, we compare our substantive findings to those using DW-NOMINATE, as this is the other dynamic model that produces point-estimates that are comparable across time. **Figure A3** shows the correlation tables for the DW-NOMINATE scores for Toronto. The first-dimension table looks remarkably similar to the one in the manuscript for our procedure's result. Thus, regardless of whether we use our solution or DW-NOMINATE, we would come to the same conclusion regarding the nature of the cleavages motivating legislator voting in Toronto. The second dimension fails to uncover much of anything meaningfully different from the first dimension, however. Our solution therefore provides a more nuanced substantive understanding of the secondary dimension.

In Chicago, the situation is quite different. Four variables are identified as significant correlates of the first dimension using the same thresholds as in the manuscript. These identify similar characteristics as in the manuscript, but with considerably less nuance. Only the percentage of Black residents is a significant correlate of the second dimension. Our procedure produces predictions that are just as good as the DW-NOMINATE solution, but allows us to identify a set of ward-level demographic correlates that illuminate the nature and important cleavages of municipal politics in Chicago.

Figure A3: DW-NOMINATE Correlation Tables

Toronto

	Pooled	1998-2001	2001-2003	2004-2006	2006-2010	2010-2014	2014-2018	2018-2022	
% car commute	0.62	0.60	0.62	0.52	0.65	0.71	0.64	0.70	place
% married	0.54	0.66	0.63	0.45	0.56	0.59	0.43	0.49	lifestyle
% Detached housing	0.46	0.46	0.39	0.36	0.53	0.54	0.45	0.58	place
% Age 0–19	0.43	0.47	0.42	0.40	0.40	0.51	0.39	0.51	lifestyle
% own home	0.38	0.59	0.32	0.27	0.45	0.44	0.32	0.41	status
% Age 65+	0.33		0.27	0.29	0.41	0.50	0.49	0.41	lifestyle
% Pentacostal religion	0.30	0.31	0.32	0.21	0.22	0.37	0.34	0.45	group
% blue-collar occupations	0.30		0.27	0.27		0.40	0.43	0.57	status
% immigrants	0.27		0.39	0.23		0.36	0.31	0.37	group
% Muslim	0.26	0.37	0.40			0.29	0.22	0.34	group
% Sikh	0.25		0.23	0.21	0.23	0.31	0.38	0.30	group
% Jamaican origin	0.22	0.24	0.23			0.23	0.23	0.35	group
% High school or less		0.48	0.47	0.30	0.24	0.46	0.44	0.64	status
% ed., law, soc. sci., and gov't occupations		-0.30	-0.28	-0.28	-0.21	-0.43	-0.35	-0.38	status
% non-immigrants	-0.23		-0.37	-0.21		-0.31	-0.24	-0.27	group
% British origin	-0.25		-0.22	-0.22		-0.37	-0.37	-0.40	group
% divorced	-0.27	-0.45	-0.44	-0.21	-0.29	-0.23			lifestyle
% rent	-0.38	-0.58	-0.32	-0.27	-0.45	-0.43	-0.32	-0.41	status
% Age 20-29	-0.42	-0.47	-0.37	-0.43	-0.49	-0.48	-0.38	-0.47	lifestyle
% No religion	-0.44	-0.41	-0.45	-0.30	-0.40	-0.53	-0.50	-0.58	group
% transit commute	-0.49	-0.46	-0.50	-0.34	-0.56	-0.57	-0.54	-0.54	place
Population Density	-0.51	-0.55	-0.43	-0.40	-0.53	-0.65	-0.52	-0.57	place
% walk/bike commute	-0.52	-0.55	-0.56	-0.48	-0.51	-0.56	-0.47	-0.55	place
% single, never married	-0.55	-0.63	-0.62	-0.51	-0.60	-0.57	-0.43	-0.50	lifestyle
% arts occupation	-0.55	-0.55	-0.62	-0.46	-0.48	-0.61	-0.52	-0.62	status

A) First Dimension

	Pooled	1998-2001	2001-2003	2004-2006	2006-2010	2010-2014	2014-2018	2018-2022	
% Catholic	0.23	0.26	0.40			0.30			group
% Portuguese origin	0.21	0.40	0.51						group
% Sikh					0.22	0.26		0.21	group
% Detached housing		-0.30				0.33		0.28	place
% married		-0.25				0.30		0.28	lifestyle
% walk/bike commute		0.27	0.26			-0.22		-0.22	place
% car commute		-0.33				0.35		0.22	place
% transit commute		0.29		-0.26	-0.23	-0.39			place
% single, never married		0.24				-0.33	-0.21	-0.26	lifestyle
Population Density		0.32				-0.38		-0.22	place
% Age 20-29		0.30	0.26			-0.28	-0.26	-0.21	lifestyle

B) Second Dimension

Chicago

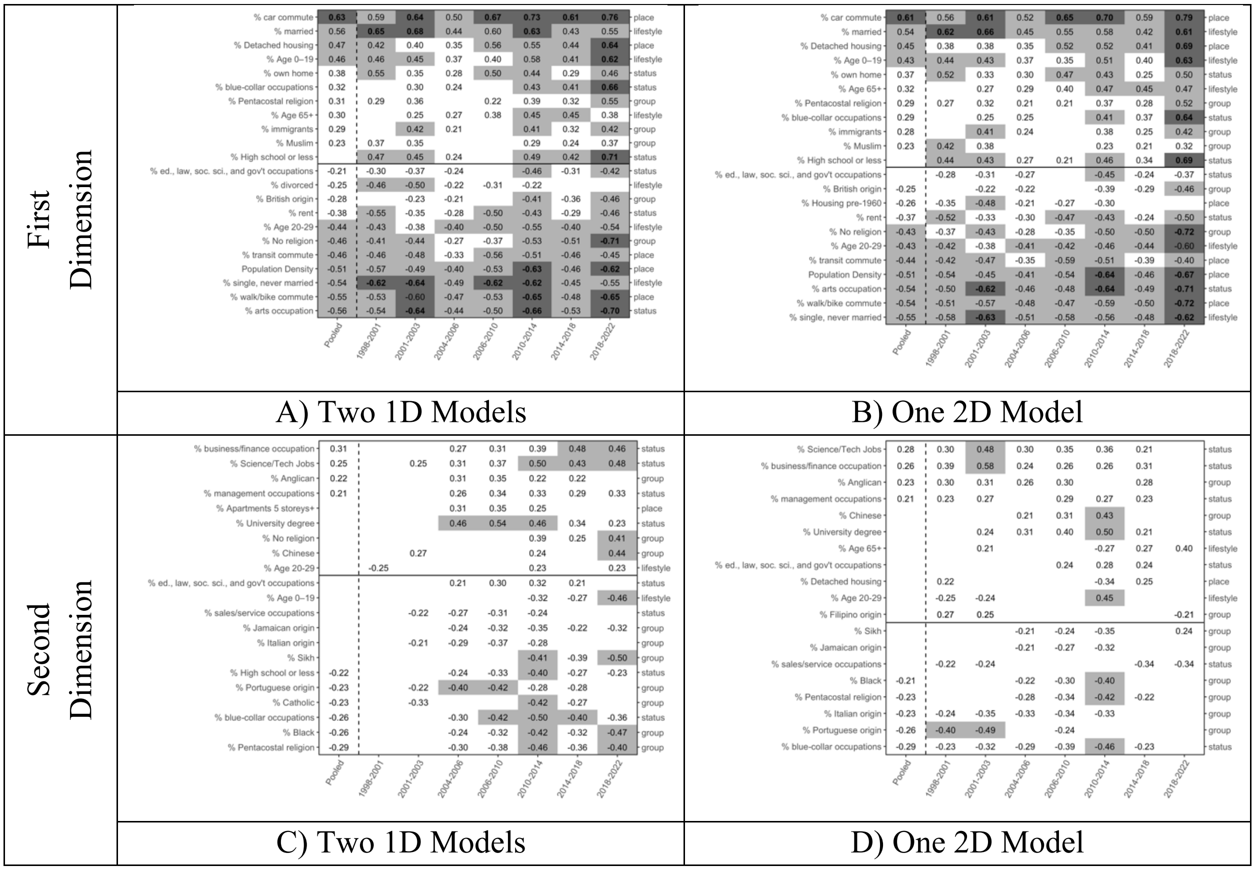
	Pooled	1979-1983	1983-1987	1987-1987	1987-1989	1989-1991	1991-1995	1995-1999	
% car commute	0.40	0.47	0.40	0.52	0.52	0.51	0.51	0.47	place
% Apartments	-0.27		-0.23	-0.38	-0.37	-0.33	-0.28		place
% rent	-0.28	-0.38	-0.30	-0.23	-0.22	-0.21		-0.27	status
Transitscore	-0.34	-0.31	-0.31	-0.41	-0.39	-0.37	-0.33	-0.29	place

C) First Dimension

Diagnostics. One final question worth indulging here is whether the findings we obtain are a function of our estimation method (i.e., the less-constrained dynamic IRT) or a function of pre-curating the data so that the different models can identify and estimate substantively different dimensions. To provide some evidence, we engaged in two different kinds of analyses. First, we estimated two one-dimensional DW-NOMINATE models based on the input data for the two dimensions from our model. Second, we estimate a single two-dimensional DW-NOMINATE model that uses our procedure's data from both dimensions as input. In both cases, we orthogonalize the result in the same fashion that our procedure does. The APRE results are very similar to the ones we obtained above (0.32 for the two one-dimensional models and 0.34 for the two-dimensional model). More importantly, we find that the substantive dimensions uncovered by the model are substantively quite similar to the ones uncovered by our procedure.

Figure A4 shows these results for Toronto. Scrutinizing the correlations, it is clear that the estimated dimensions are quite similar to the ones presented in the manuscript. If we attend to the truly independent sources of variation in the data, we can uncover similar results regardless of the estimation method (at least in choosing between our less-constrained dynamic IRT and the DW-NOMINATE models).

Figure A4: Correlation Tables using Curated Variables used in our Procedure



The results for Chicago are less encouraging. First, DW-NOMINATE will estimate the model with all the data, but will not estimate the model with the curated variables used in our procedure – there are too few observations in periods 5, 7 and 8 for the model to converge. DW-NOMINATE operates by estimating the W-NOMINATE measures for each period and then modifies the results to make them comparable over time. Thus, if the W-NOMINATE model will not estimate for any particular period, then DW-NOMINATE is stymied. Our approach uses a one-shot dynamic model where ideal points in periods of sparse data will be a compromise between the previous period's ideal point and the current period estimated value, though the balance of the compromise will favour the previous period's value. In periods of abundant data, the compromise will be weighted more in favour of the current period estimation. In statistical terms, there is shrinkage in the direction of the previous period's estimate where the magnitude of the shrinkage is a function of the precision of this period's estimate relative to the precision of the previous period's estimate. We find no meaningful correlations with the ward profile data when we estimate the DW-NOMINATE model for all the Chicago data simultaneously and orthogonalize as above.

From this analysis, we conclude that in a wide variety of situations, the solution we propose produces reasonable estimates that can help illuminate the nature of electoral competition. In some cases, DW-NOMINATE with a similar post-estimation orthogonalization can perform similarly well (particularly in abundant data situations), but our experience with the Chicago data suggests that there are important limitations to the DW-NOMINATE procedure's ability to identify distinct, important sources of underlying variation.

Appendix B: Ward profile data

The Chicago ward profiles were assembled from census tract–level data from the 1970–2010 decennial census as disseminated in the Longitudinal Tract Database (Logan, Xu, and J. 2014) with additional variables added from the National Historical Geographic Information System (McMaster, Lindberg, and Van Riper 2003). The Toronto ward profiles were created from census tract–level data from the 1996–2016 quinquennial census and harmonized to common spatial units using the Canadian Longitudinal Tract Database (Allen and Taylor 2018). Tract-level data were apportioned to wards by decomposing them into their constituent census blocks, weighted by block population, and then summing the block values to their wards. Census data are assigned to ward schemes and terms of office based on proximate year of collection. These assignments are shown in **Table B1**. As the American and Canadian census bureaux collect different information, the same variables are not available for each city. However, we have taken care to ensure that each social concept – race, ethnic origin, occupation, housing type, and so on – is represented in both cities.

For Chicago, tract-level indicators of amenity accessibility were drawn from the Walk Score service and aggregated to ward boundaries. For any geographic location in the United States, Walk Score calculates indexes representing accessibility to amenities by walking, transit, and cycling, accounting for street connectivity, topography, and the capacity and frequency of the transit network. Researchers have correlated higher walk scores with the propensity to walk, lower rates of obesity, lower mortgage default rates, and higher residential and commercial property values (WalkScore 2021). Ward-level walk and transit scores were calculated by population-weighted averaging scores retrieved for ten randomly selected locations in each census tract.

The available variables used for each city, sorted by type (group, status, place, and lifestyle), are summarized in **Tables B2 and B3**.

Table B1: Data sources and temporal assignments

Chicago

Term of office	*Ward scheme*	*Data year*
1979–83	1970	Census 1980
1983–87	1981	Census 1980
1987–91	1986	Census 1980
1991–95	1986	Census 1980
1995–99	1992	Census 1990
1999–2003	1998	Census 2000
2003–07	2002	Census 2000
2007–11	2002	Census 2010
2011–15	2002	Census 2010
2015–19	2012	Census 2010

Toronto

Term of office	*Ward scheme*	*Data year*
1998–2000	1998 (28 wards)	Census 1996
2000–03	2000 (44 wards)	Census 2001
2003–06	2000 (44 wards)	Census 2001
2006–10	2000 (44 wards)	Census 2006
2010–14	2000 (44 wards)	National Household Survey 2011
2014–18	2000 (44 wards)	Census 2016

Table B2: Ward characteristic variables – Chicago

Group variables	**Status variables**
Ethnicity/Origins	*Employment*
% Cuban	% Unemployed
% German	
% Irish	*Housing Cost*
% Italian	Housing Cost Index*
% Mexican	
% Polish	*Educational Attainment*
% Puerto Rican	% High School or Less
% Russian	% University Degree
% Scandinavian	
	Income
Immigration Status	$ Median Household Income (Standardized)
% Foreign-born	
	Occupation
Race	% Arts and Recreation
% Black (nonhispanic)	% Blue Collar
% Chinese	% Manufacturing
% Hawaiian and Pacific Islander	% Management
% Latino	% Professionals
% South Asian	% Sales and Services
% White (nonhispanic)	
	Housing Tenure
	% Own Home
	% Rent
Place Variables	**Lifestyle Variables**
Amenity Profile	*Age Cohorts*
Transitscore	% Young Adults 18-25
Walkscore	% Seniors 60+
Bikescore	
	Marital Status
Housing Type and Age	% Divorced
% Apartments	% Married
% Housing More than 30 Years Old	
% Housing Vacant	
Mobility	
% Commute to Work by Car	
% Commute to Work by Transit	
% Commute to Work by Walk, Bike	
Distance from City Hall	
Population Density (per km2, Standardized)	

* Housing cost index calculated using Chronbach's alpha from standardized owned dwelling value and standardized rent cost (α = .95 in Chicago).

Table B3: Ward characteristic variables – Toronto

Group variables	**Status variables**
Ethnicity/Origin % British % Chinese % Filipino % Italian % Jamaican % Portuguese % Sri Lankan/Tamil *Immigration Status* % Foreign-Born *Race (Visible Minority Status)* % Arab % Black % Chinese % Non-White % South Asian % White *Religion* % Anglican % Catholic % Hindu % Jewish % Muslim % No religion (Atheist/Irreligious) % Pentacostal % Sikh % United Church of Canada	*Housing Cost* $ Average Home Value (Standardized) $ Monthly Rent (Standardized) *Educational Attainment* % High School or Less % University Degree *Income* $ Average Household Income (Standardized) *Occupation* % Arts and Recreation % Blue Collar (Manufacturing, Trades, Construction) % Business, Finance, and Administration % Education, Law, Social Science, Government % Management % Professionals % Sales and Services % Science and Technology *Housing Tenure* % Own Home % Rent
Place variables	**Lifestyle variables**
Housing Type and Age % Apartments % Single-Detached Housing % Housing Built in 1960 or earlier *Mobility* % Commute to Work by Car % Commute to Work by Transit % Commute to Work by Walk, Bike Population Density (per km2, Standardized)	*Age Cohorts* % Children 0–19 % Young Adults 20–29 % Seniors 65+ *Marital Status* % Married % Single, Never Married % Divorced

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