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Why Do Liberals Drink Lattes?[1]

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Popular accounts of "lifestyle politics" and "culture wars" suggest that political and ideological divisions extend also to leisure activities, consumption, aesthetic taste, and personal morality. Drawing on a total of 22,572 pairwise correlations from the General Social Survey (1972–2010), the authors provide comprehensive empirical support for the anecdotal accounts. Moreover, most ideological differences in lifestyle cannot be explained by demographic covariates alone. The authors propose a surprisingly simple solution to the puzzle of lifestyle politics. Computational experiments show how the self-reinforcing dynamics of homophily and influence dramatically amplify even very small elective affinities between lifestyle and ideology, producing a stereotypical world of "latte liberals" and "bird-hunting conservatives" much like the one in which we live.

> In content, status honor is normally expressed by the fact that above all else a specific *style of life* can be expected from all those who wish to belong to the circle. . . . As soon as there is not a mere individual and socially irrelevant imitation of another style of life, but an agreed-upon communal action of this closing character, the "status" development is under way.
>
> (Max Weber, *Class, Status, and Party*)

> What do I think? Well, I think Howard Dean should take his tax-hiking, government-expanding, latte-drinking, sushi-eating, Volvo-driving, *New York Times*–reading, body-piercing, Hollywood-loving left wing freak show back to Vermont where it belongs.
>
> (2004 Club for Growth TV advertisement)

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THE PUZZLE OF LIFESTYLE POLITICS

A century of research has documented the clustering of populations into internally homogeneous "lifestyle enclaves" (Park 1915; Garreau 1981; Bellah et al. 1985; Weiss 1988; Lamont et al. 1996; Fischer and Hout 2006; Bishop 2008; Fischer and Mattson 2009). The term "lifestyle" has been criticized as superficial and ambiguous—a conceptual "grab bag" that lacks a theoretical foundation (Zablocki and Kanter 1976). Yet empirical studies provide consistent evidence of clustering across seemingly eclectic "lifestyle" dimensions, ranging from hot-beverage choices to musical tastes to a preference for vowels (Weiss 1988; Bishop 2008; Hall-Lew, Coppock, and Starr 2010; Labov 2010). On the basis of a comprehensive review of the literature on political and cultural polarization "by cultural sociologists, scholars of consumption, media analysts, and journalistic accounts," Fischer and Mattson (2009, p. 446) conclude that "the number of new, discrete, and separated social worlds increased between 1970 and 2005." Simply put, we are increasingly likely to find our local communities and social networks populated by individuals with similar aesthetic tastes, leisure activities, consumer preferences, moral practices, and ways of life.

These disparate cultural profiles often differ as well in political ideology and behavior, as noted in studies of "lifestyle politics" and "culture wars" (Miller 1985; Giddens 1991; Hunter 1991; Bennett 1998; Chaney 2002). Although consumer preferences rarely engender heated debate, lifestyle can become controversial when it comes to be associated with or symbolic of membership in a group with an identifiable political or ideological profile. A classic example is Gusfield's (1963) study of the temperance movement. Gusfield argues that temperance was not about a preference for nonalcoholic beverages but a symbol of white Anglo-Saxon Protestant nativism. The 1960s counterculture is a more recent example in which long hair, acid rock, and paisley dresses came to be associated with opposition to the Vietnam War. Two decades later, following a pilgrimage to India, Steve Jobs helped Apple computers come to be identified with the counterculture, in contrast to the button-down world of IBM. Preferences for hunting, body modification, country-western music, hip-hop, and even fast-food consumption have acquired normative overtones as these preferences

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have come to be associated with politically and ideologically distinctive communities.

Recent research by Baldassarri and Gelman (2008) provides compelling evidence from opinion surveys demonstrating strong and increasing political and ideological alignment on what they call "moral issues." When lifestyle choices such as whom to marry, child-rearing responsibilities, and when and whether to pray or have children come to be widely regarded as expressions of group membership, they can become divisive. These "hot-button" issues—including gay rights, gender roles, school prayer, and abortion—differ from consumer tastes in focusing on normative rather than aesthetic preferences and, accordingly, in the level of passion they engender. For example, after fast-food chain Chick-Fil-A announced opposition to same-sex marriage, purchasing a sandwich suddenly became an expression of political identity as same-sex marriage opponents flocked in droves to take part in "Chick-Fil-A Appreciation Days" while proponents picketed outside. Recent evidence suggests that even the choice of partisan marital partners has become divisive: Democrats and Republicans increasingly report that they would be upset if their children married someone of the opposite party (Iyengar, Sood, and Lelkes 2012).

In short, the puzzle of lifestyle politics compounds the curious formation of cultural enclaves among seemingly unrelated preferences. Why should liberals and conservatives differ systematically on lifestyle dimensions that have no apparent substantive relevance to political ideology? What are the social mechanisms that could produce a world of "latte liberals" and "bird-hunting conservatives"?

We begin by presenting survey data that support anecdotal accounts of lifestyle politics. To that end, we replicate Baldassarri and Gelman's (2008) analysis of political and ideological alignment using different data (the General Social Survey [GSS]) but then go beyond the explicitly political hot-button issues in their study (e.g., abortion rights and same-sex marriage) to demonstrate the ideological clustering of seemingly nonideological lifestyle items such as belief in astrology, art consumption, and leisure activities.[2] These questions of taste are rarely contested in the back-and-forth of ideological debate between political parties or in commentary by pundits. Yet our analysis finds that these aesthetic lifestyle dimensions are also surprisingly correlated, not only with other lifestyle preferences but also with measures of political and ideological affiliation. In some cases, the correlations are as strong as or stronger than partisan differences over hotly debated social issues.

[2] Unfortunately, hot-beverage preferences are not measured in the GSS, but there are numerous items on leisure activities and consumption.

We then pursue possible solutions to the puzzle of lifestyle politics, beginning with a review and critique of the prevailing explanatory strategy that attributes opinions and lifestyle preferences to the formative effects of causally prior demographic, socioeconomic, and cultural backgrounds. These explanations are intuitively compelling and often well documented with data from opinion surveys, but they are constrained by the absence of network relations in the random samples on which most surveys are based. Decades of survey-based opinion research have lacked the relational data with which to rule out network autocorrelation as an alternative explanation. Network autocorrelation refers to the dependence of individual attributes on those of network neighbors and is the relational equivalent of temporal and spatial autocorrelation.

In response, we propose a highly parsimonious alternative based on a formalization of McPherson's (2004) theory of an "ecology of affiliation." We build on and extend previous models of interpersonal influence and homophilous selection in the dynamics of political and cultural differentiation (Carley 1991; Axelrod 1997; Mark 2003; Centola et al. 2007; Mäs, Flache, and Helbing 2010; Flache and Macy 2011*a*) and polarization (Macy et al. 2003; Kitts 2006; Baldassarri and Bearman 2007; Macy and Flache 2009). We use an agent-based computational model (Macy and Willer 2002) to identify conditions in which the self-reinforcing dynamics of homophily and social influence can carve deep cultural canyons in a demographic landscape. When we analyzed the data generated by the simulations, treating each individual as an independent observation, the results were consistent with the conclusion that lifestyle correlations are partly explained by demographic background, even though the correlations were largely (though not entirely) an artifact of network autocorrelation.

Finally, we conclude by discussing the implications of network autocorrelation for studies of cultural polarization, social influence, and political behavior. These implications include troubling questions about the use of random samples to study opinion formation in a population whose members share their opinions with others.

To be clear, we do not directly answer the riddle posed in our title, nor do we offer explanations for any other specific lifestyle correlations. We use the popular image of the "latte liberal" only to illustrate the more general conundrum of lifestyle politics that motivates our proposed explanation. We present a relational model of path-dependent opinion dynamics and group formation that points to the need for historical case studies for understanding the emergent self-organization of specific instantiations of the broader pattern, but the model also cautions against elevating that understanding to the level of general theory.

LIFESTYLE POLITICS IN THE GENERAL SOCIAL SURVEY

Using data from the American National Election Study, Baldassarri and Gelman (2008) document a broad pattern of zero-order correlations between attitudes on what they characterize as "moral issues" and self-reported political ideology. We replicate their analysis using different data and with an important new twist: In addition to the hot-button political issues found in popular and scholarly accounts of lifestyle politics (e.g., abortion and same-sex marriage), we broaden the scope to also include leisure activities, consumer preferences, aesthetic tastes, and personal convictions for which correlation with political views is considerably more puzzling.[3]

In this section, we first report results showing that anecdotal accounts of lifestyle clustering are strongly supported by survey data from the GSS cumulative file (1972–2010; Smith, Marsden, and Hout 2010). We then use the cumulative file to provide a more rigorous systematic answer to the empirical question posed by media and popular depictions of "latte liberals": to what extent are lifestyle preferences correlated with political and ideological alignment?

Lifestyle Clustering

The GSS is a large and nationally representative survey of American demographics, behaviors, attitudes, and opinions. The survey was administered 28 times between 1972 and 2010 by the National Opinion Research Center at the University of Chicago. We identified 216 GSS items that fit with a definition of lifestyle as pertaining to "shared values or tastes as reflected primarily in consumption patterns but applicable also to the evaluation of intangible and/or public goods" (Zablocki and Kanter 1976, p. 270). These include questions on preferred musical genres, attitudes toward art, leisure activities, and consumer preferences. They also include normative attitudes and beliefs that are likely to influence one's own behavior but are rarely imposed on others, for example, attitudes toward marriage, children, divorce, altruism, and the definition of "right" and "wrong." In addition, we examine ideological differences on hot-button issues that form the battlegrounds of the "culture wars," including same-sex marriage, abortion, gender roles, and school prayer.[4] Where necessary,

[3] Baldassarri and Gelman (2008) also focus primarily on temporal shifts in the pattern of correlation. Our interest centers on the cumulative associations, and we leave the changes over time for future research.

[4] The selected items are based on examination of the cumulative GSS codebook available at http://www.norc.org/GSS+Website/Documentation/. We excluded redundant

items were recoded in order to produce binary, metric, or ordinal scales. Alternative phrasings of the same questions were consolidated. Nonmetric nominal items (e.g., religious affiliation) were not used. (Our online supplement provides a more detailed description of each of the 216 lifestyle items and how they were coded.)

Using these 216 lifestyle measures, we generated year-specific correlations for every observed instance in which two items appeared in the same edition of the GSS. This exercise produced a new data set with 21,625 correlations nested within 10,180 unique pairs of lifestyle items. Following Baldassarri and Gelman (2008), we used a linear mixed model to correct possible biases caused by variation in the number of GSS editions for which responses are available. While the items appear in surveys ranging from 1972 to 2010, some items appear in just one survey and others in nearly all. An unweighted average over a given vector of year-specific pairwise correlations can be heavily biased toward items that were asked in many GSS editions, typically because they resonate with the political zeitgeist. To produce estimates of average year-specific correlation that account for the variation between item pairs appearing at different time points and with different frequency, we estimated a mixed-effects model with varying intercepts and slopes (Gelman and Hill 2007). We fit the model using Markov chain Monte Carlo (MCMC) techniques.[5]

Formally, we index pairs of lifestyle items by j. The magnitude of the correlation between the items in pair j for year t is given by

$$|r|_{j,t} = \alpha_j + \beta_j t + e_{j,t}, \tag{1}$$

where α_j is the intercept that varies by each item pair j, β_j is a pair-specific time trend, and $e_{j,t}$ is the residual deviation from the intercept for year-specific instantiations of the item pair.[6]

In addition to estimating this model for all zero-order correlations between lifestyle items, we also obtained corresponding partial correlations,

lifestyle measures so as not to inflate the observed proportion of items that were correlated with ideology. For example, the GSS includes several items about hunting, but we included only the most general measure.

[5] We use the "MCMCglmm" package in R (Hadfield 2010). We also replicated the analysis using the "xtmixed" package in Stata and found that the 2010 point estimates were virtually identical ($r = .999$). Given the complexity of the model's error structure, we prefer the MCMC approach because it provides a more straightforward avenue for quantifying the uncertainty in these point estimates.

[6] Of the 216 items, 107 appeared in just one edition of the survey. We reestimated the mixed-effects models excluding these observations and found larger estimated intercepts than when they were included. Accordingly, we present conservative results that incorporate the broadest range of GSS items.

controlling for 10 measures of demographic and socioeconomic background: marital status (married, widowed, divorced, separated, or never married); parenthood (parent vs. nonparent); religious denomination (Protestant, Catholic, Jewish, other denomination, or none); region (New England, Mid-Atlantic, East Northern Central, West Northern Central, South Atlantic, East Southern Central, West Southern Central, Mountain, or Pacific); highest educational degree (less than high school, high school, junior college, bachelor's degree, or graduate degree); logged family income in constant dollars; logged size of place (measured in thousands of residents); sex; birth cohort; and race (white, black, or other).

Figure 1 plots two sets of results from the linear mixed models. For both zero-order and partial correlations between pairs of lifestyle items, figure 1 shows the univariate kernel density curve of year-specific correlation magnitudes fitted by the model. In addition, the vertical reference lines show model averages capturing the expected magnitude of correlation for an item pair. For consistency and recency, we graph estimates for the expected correlation magnitudes in 2010.[7] An older reference point would shift the mean estimates slightly, but not enough to change the substantive story (see table A1 in the appendix). In the remainder of this section, we present estimates for a number of selected individual correlations. Our purpose is to illustrate both the intuitive and surprising results of our data-mining exercise. For each item pair, we report model-estimated values corresponding to the predicted zero-order ($|\hat{r}_0|$) and partial ($|\hat{r}\cdot|$) correlation magnitude in 2010. Using the MCMC-generated posterior density intervals, we test each 2010 point estimate against the null hypothesis of zero correlation between items and use asterisks to signify the likelihood of this null result (* for <.05; ** for <.01; *** for <.001).

Unsurprisingly, correlation tends to be strongest between items that tend to be colocated on lists of sinful indulgences. For example, respondents who agreed that "premarital sex is always wrong" also reported spending fewer of their evenings in bars or taverns ($|\hat{r}_0| = .30^{***}$; $|\hat{r}\cdot| = .19^{***}$) and were less likely to say that they "ever have occasion to use any alcoholic beverages such as liquor, wine, or beer" ($|\hat{r}_0| = .37^{***}$; $|\hat{r}\cdot| = .26^{***}$). Less obvious is the strong negative correlation between disapproving of premarital sex and liking contemporary rock ($|\hat{r}_0| = .30^{***}$; $|\hat{r}\cdot| = .16^{***}$) or the positive correlation between going to a bar and agreeing with the statement that "homosexual couples should have the right to marry one another" ($|\hat{r}_0| = .25^{***}$; $|\hat{r}\cdot| = .15^{***}$). Attitudes toward homosexuality are also strongly aligned with a number of other lifestyle preferences. For

[7] Forty-five of the 216 items appeared most recently in the 2010 GSS, and 45% of the items have appeared at least once since 2000.

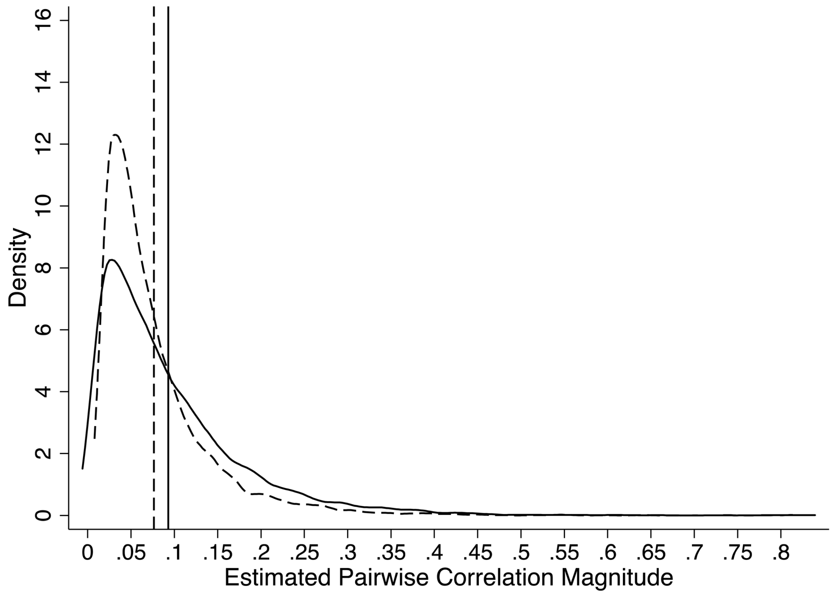

FIG. 1.—Magnitude of zero-order and partial correlation among GSS lifestyle items. Graphs plot Epanechnikov kernel density functions for both zero-order (solid lines) and partial (dashed lines) correlation magnitudes estimated from the mixed-effects model (see table A1 in the appendix). One value is plotted for each of the 10,180 item pairs. Time is set to 2010 to facilitate comparison across item pairs. The solid vertical reference line gives the mean predicted zero-order correlation magnitude across all item pairs in 2010 and the dashed vertical reference line gives the same value for partial correlation magnitude.

example, those who like reggae music ($|\hat{r}_0| = .25^{***}$; $|\hat{r}\cdot| = .15^{***}$) and read fiction ($|\hat{r}_0| = .17^{***}$; $|\hat{r}\cdot| = .12^{***}$) were far less likely to agree with the statement that "sexual relations between two adults of the same sex are always wrong."

Lifestyle Politics

The puzzling association between seemingly unrelated lifestyle preferences is further compounded by the tendency for lifestyles to be correlated with political ideology. For each edition of the GSS, we measured the pairwise zero-order correlation between each lifestyle item observed in that survey and self-reported political ideology (measured on a seven-point scale from "extremely conservative" to "extremely liberal"). In other words, the lifestyle-item pairs indexed by j in equation (1) instead become lifestyle-ideology pairs. Figure 2 graphically depicts the results of this second set

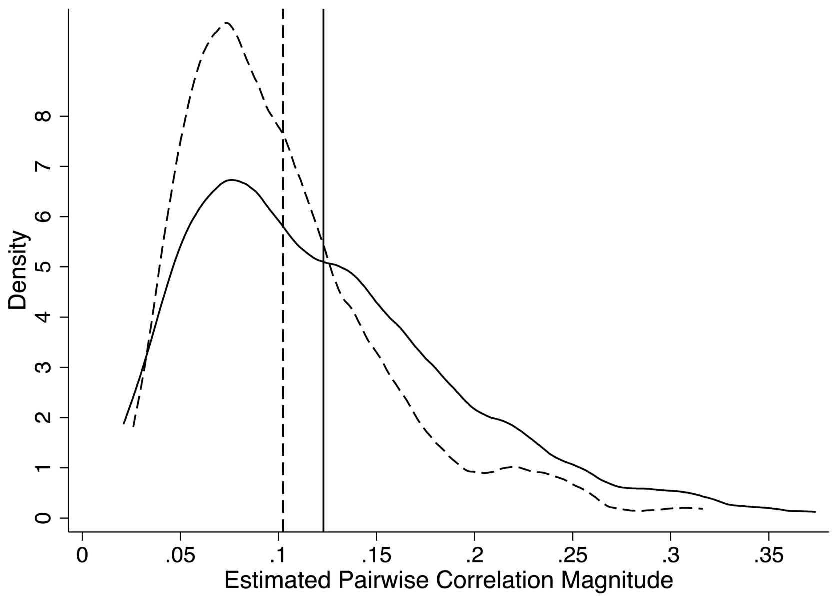

FIG. 2.—Magnitude of zero-order and partial correlation between GSS lifestyle items and ideological identity. Graphs plot Epanechnikov kernel density functions for both zero-order (solid lines) and partial (dashed lines) correlation magnitudes estimated from the mixed-effects model (see table A2 in the appendix). One value is plotted for each of the 216 item pairs. Time is set to 2010 to facilitate comparison across item pairs. The solid vertical reference line gives the mean predicted zero-order correlation magnitude across all item pairs in 2010 and the dashed vertical reference line gives the same value for partial correlation magnitude.

of linear mixed models.[8] The expected zero-order correlation between one of the 216 lifestyle items and liberal-conservative ideology in 2010 is .12 in absolute value. Again, we also tested the extent to which these zero-order correlations reflect the joint effects of causal priors, namely, demographics and socioeconomic position: the expected partial correlation in 2010 is .10 in absolute value.

Among the 216 MCMC estimates of 2010-adjusted correlation magnitude, about 58% were statistically significant at the .05 level. The average magnitude of .12 and increases to .16 in the subset of statistically significant correlations. For 50 of the 216 lifestyle-politics pairs, controlling for indi-

[8] Tabular versions of these results are given in table A2 in the appendix, where we also show the results of similar models replacing political ideology with party identification. We find qualitatively similar results, though correlations between lifestyle and party identification are generally weaker than those with ideology.

vidual background actually increases the magnitude of the correlation, indicating a suppressor effect. Among the remaining 166 pairs, correlations decline on average by about 21%. The partial correlation is no longer significant at the .05 level for 26 of these, which shows that the zero-order correlation is causally spurious. However, this does not mean that the observed association is necessarily inconsequential, unimportant, or uninteresting. For example, we found that the association between liberalism and preference for heavy metal music is reduced by about 43% (from .16 to .09) when controlling for sociodemographic background. This indicates that political ideology does not affect heavy metal appreciation (or vice versa), but it does not mean that the formation of sociodemographically differentiated political-lifestyle enclaves is any less puzzling.

Clearly, statistical significance does not necessarily correspond to theoretical significance. An average .12 zero-order correlation between lifestyle items and political ideology in 2010 means that many of these items are weak predictors of whether a person is liberal or conservative. However, the average obscures individual correlations, many of which are large enough to warrant explanatory efforts. Particular instantiations of the broader pattern of correlation range from empirically familiar to highly surprising. Consistent with customary stereotypes, liberals were less likely to report praying often ($|\hat{r}_0| = .21^{***}$; $|\hat{r}\cdot| = .14^{***}$) or owning any guns ($|\hat{r}_0| = .17^{***}$; $|\hat{r}\cdot| = .07^{***}$) and much more likely to feel that "animals should have the same moral rights that human beings do" ($|\hat{r}_0| = .18^{***}$; $|\hat{r}\cdot| = .12^{**}$).

Musical tastes are frequently correlated with liberalism and conservatism, as shown when respondents were asked to rank a series of musical genres on a scale from "dislike very much" to "like very much." The resulting correlations between ideology and musical preferences reflect an omnivorous liberal affinity for New Age ($|\hat{r}_0| = .18^{***}$; $|\hat{r}\cdot| = .12^{*}$), blues/rhythm and blues ($|\hat{r}_0| = .16^{**}$; $|\hat{r}\cdot| = .12^{*}$), reggae ($|\hat{r}_0| = .16^{**}$; $|\hat{r}\cdot| = .10$), and jazz ($|\hat{r}_0| = .15^{**}$; $|\hat{r}\cdot| = .10$). Liberals were more likely to view rock music as a positive influence on children ($|\hat{r}_0| = .15^{*}$; $|\hat{r}\cdot| = .11^{*}$) and to have attended a live performance of popular music "like rock, country, or rap" in the previous 12 months ($|\hat{r}_0| = .13^{***}$; $|\hat{r}\cdot| = .10^{**}$). Partisans further differ in their assessments of nonmusical artistic endeavors: liberals are less apt than conservatives to agree with the statement that "modern painting is just slapped on; a child could do it" ($|\hat{r}_0| = .15^{***}$; $|\hat{r}\cdot| = .13^{***}$).

Conservatives are more likely to agree that "we trust too much in science and not enough in religious faith" ($|\hat{r}_0| = .20^{***}$; $|\hat{r}\cdot| = .15^{***}$). This would seem to suggest that conservatives are also more inclined to believe in metaphysical explanations, magic, and superstition. Surprisingly, we found the opposite. The liberal attitude does not appear to be grounded in a consistently scientific worldview; on the contrary, liberals are more likely than conservatives to have read horoscopes or personal astrology reports ($|\hat{r}_0| = .12^*$; $|\hat{r}\cdot| = .11^{***}$) and to believe in the "supernatural powers of deceased ancestors" ($|\hat{r}_0| = .12^{***}$; $|\hat{r}\cdot| = .10^*$).

Consistent with scholars' and pundits' characterizations of "culture wars," much stronger pairwise correlations are observed for hot-button issues such as homosexuality, abortion, sexual mores, legal marijuana, and euthanasia. For example, the zero-order correlation between liberal ideology and approval of same-sex marriage (see above for exact wording) is large ($|\hat{r}_0| = .37^{***}$) and declines little when adjusted for demographic and socioeconomic covariates ($|\hat{r}\cdot| = .32^{***}$). Liberals and conservatives also differ markedly in their moral judgments. Conservatives, for example, are more likely to agree that "it is sometimes necessary to discipline a child with a good, hard spanking" ($|\hat{r}_0| = .15^{***}$; $|\hat{r}\cdot| = .12^{***}$). They are also more likely to feel that "a marriage without children is not fully complete" ($|\hat{r}_0| = .16^{**}$; $|\hat{r}\cdot| = .13^*$). Liberals are overwhelmingly more likely than conservatives to agree that "it's a good idea for a couple who intend to get married to live together first" ($|\hat{r}_0| = .32^{***}$; $|\hat{r}\cdot| = .26^{***}$). They are also more likely than conservatives to feel that "when a marriage is troubled and unhappy, it is generally better" for the husband ($|\hat{r}_0| = .22^{**}$; $|\hat{r}\cdot| = .16^{**}$), wife ($|\hat{r}_0| = .19^{**}$; $|\hat{r}\cdot| = .14^*$), and children ($|\hat{r}_0| = .19^{**}$; $|\hat{r}\cdot| = .14^*$) if the couple gets divorced rather than staying together. Perhaps most telling of all, liberals are much more likely to agree with the statement that "right and wrong are not usually a simple matter of black and white; there are many shades of gray" ($|\hat{r}_0| = .20^{***}$; $|\hat{r}\cdot| = .17^{***}$).

In total, 54% of the 216 lifestyle-ideology pairs produce estimated 2010 correlations of magnitude .10 or higher. Thirty-four percent meet a higher threshold of .15, and 89% of these remain at or above .10 in magnitude with the inclusion of demographic and socioeconomic covariates. The list of items that meet both thresholds includes not only hot-button political issues such as homosexuality, abortion, and capital punishment but also attitudes toward modern art, music, science, and animal rights. It is not difficult to find published papers that report effect sizes of similar magnitude; the important finding is not that all the correlations are large, but that many are large enough to warrant a systematic effort at explanation. It is to that task that we now turn.

EXPLAINING THE PUZZLE

A standard approach to causal analysis in opinion research attributes the more changeable side of a correlation to the less changeable (Davis 1985). Applied to lifestyle politics, horoscope reading, for example, is more likely to reflect a respondent's underlying political ideology than the other way around, and both lifestyle preferences and political ideology might be attributed to still more deeply rooted covariates, such as material interests, formative experiences, and fundamental cultural values, which in turn might be traced to demographic traits, such as cohort, schooling, gender, ethnicity, or socioeconomic background.

This explanatory strategy owes much to Weber's theory of the status group as a community of individuals with a shared "style of life" ([1946] 1991). While classes are defined in terms of the *production* of goods, these status groups and their associated lifestyles stem instead from patterns of *consumption*. The link between production and consumption confers an "elective affinity" between social class and membership in a status group. Veblen (1934) famously finds rampant and conspicuous consumption of leisure goods among the upper classes, whose members flaunt their lavish lifestyles in order to affirm their superior status. Lipset and Rokkan's (1967) survey-based studies of "class politics" show how life chances and experiences shape respondents' attitudes, preferences, and opinions. Bourdieu ([1979] 1984) further explains the relationship between social-structural location and lifestyle in terms of the mediating effects of the habitus, which structures the understanding and appreciation of cultural objects. Cultural tastes are thus indelibly shaped by the material conditions of existence.

More recent empirical work has pointed to broad social forces that impose wide variation across and within cultures with regard to both the importance and content of lifestyle. Increasing education and affluence are found to increase the general importance of lifestyle relative to race, geographic location, ethnicity, and class (Giddens 1991; Inglehart 1997; Clark and Hoffman-Martinot 1998). In a recent study, for example, Newman and Bartels (2011) find that individuals with higher educational attainment are more likely to make consumer decisions that reflect their political leanings. Socioeconomic location further structures lifestyle by shaping access and exposure to lifestyle objects, such as the arts (DiMaggio and Useem 1978) and information technology (DiMaggio et al. 2004). The unique lifestyle correlates of education and upper-class membership are perhaps best encapsulated in the image of the highbrow "cultural omnivore" (Peterson and Kern 1996; for a recent analysis, see Goldberg [2011]). Education may also have indirect effects on lifestyle via the effects of occupational self-direction on personality and psychological functioning, as shown by Kohn (1969) and Kohn and Schooler (1983). Work that is substantively complex, with minimal supervision and routinization, encourages incumbents to question tradition and social rank as bases for authority, to be less willing to conform to norms that lack credible justification, and to be more tolerant of ambiguity and diversity—hence the association between "class and conformity," the title of Kohn's seminal treatise.

Demographic and socioeconomic backgrounds are not the only organizing principles that researchers have investigated. The larger average GSS correlations for issues of private and public morality compared to personal tastes may reflect underlying moral divisions. Lakoff (1996) argues that political parties base their programs on differing models of the family (the Republicans' strict father versus the Democrats' nurturing mother) that are in turn deeply rooted in moral worldviews. Haidt (2012) argues for six innate "moral foundations" about which liberals and conservatives differ. Liberals are attracted primarily to the care, fairness, and liberty foundations, while conservatives favor all six, including loyalty, authority, and sanctity. Similarly, Jost, Federico, and Napier (2009) argue that conservatives and liberals differ fundamentally in their need for certainty, making dogmatism and intolerance of ambiguity typical of the political right, in contrast to liberals' openness to new experiences and cognitive complexity. For all these authors, moral and cognitive principles are deeply rooted and therefore operate much like demographics and socioeconomic background as cultural primitives that act as organizing principles in the alignment of morally relevant opinions.

The common thread among these theories is the argument that demographic and ideological organizing principles create axes of culture through the formative experiences, material interests, and moral foundations that can be statistically captured by indicators of social and cultural background. In the nearly six decades since Stouffer and colleagues developed survey instruments to measure attitudes in representative cross sections of the American public (Stouffer 1955), this explanatory strategy has yielded many high-impact papers and important contributions.

Nevertheless, our systematic sweep through the GSS shows that ideological alignment is not limited to a small number of high-profile lifestyle issues—such as abortion and same-sex marriage—that have been championed by political parties as expressions of contested moral principles. Aesthetic taste in music and art, belief in the supernatural (e.g., horoscope reading), and leisure activities (e.g., attending a concert) do not feature prominently in the back-and-forth of political debate, yet the absence of partisan contestation does not mean that these lifestyle dimensions are apolitical. On the contrary, the GSS analysis shows how the politicization of lifestyles extends to a diverse range of nonpartisan dimensions.

The large number of observed correlations invites an equally large number of theories about lifestyle differences between liberals and conservatives. Recall how liberals have greater confidence in science as well as New Age spiritualism. This inconsistency in liberal embrace of a scientific worldview illustrates a broader message that is suggested by the ideological clustering of lifestyle preferences: the empirical regularities have no obvious or apparent logical regularity. We do not observe a pattern of correlations that would suggest an underlying belief in mysticism, metaphysics, xenophobia, libertarianism, egalitarianism, or existentialism that might logically account for the differences between liberals and conservatives in lifestyle choices and preferences. Without a coherent cultural or philosophical underpinning for the observed correlations, each correlation may require a uniquely specialized explanation and a rich theoretical imagination. The search for causal priors then becomes vulnerable to "just-so stories" and overfitting. While a plausible story might be constructed to explain ideological differences on any given lifestyle dimension, an equally plausible story could usually be constructed were the differences to be reversed. For example, if conservatives are more skeptical of astrology, the reason is that astrology is regarded as sacrilegious; and if conservatives are less skeptical, the reason is that they feel less need for scientific proof.

Ironically, the empirical puzzle is often most perplexing when trying to explain opinions outside the lifestyle domain, such as affirmative action, social welfare policy, and government regulations for which ethnic, regional, and class membership have immediate relevance. Although these issues have clear economic implications, demographic and socioeconomic explanations have encountered decades of controversy and contradictory findings. Why do struggling farmers in Kansas vote for conservative Republicans (Frank 2004), while so many wealthy Connecticut suburbanites favor progressive Democrats (Gelman et al. 2008)? And why was one of the closest presidential vote in U.S. history in 2000 (with a razor-thin margin of victory at the national level) nevertheless often a landslide one way or the other at the county level (Bishop 2008)?

A RELATIONAL ALTERNATIVE

Statistical tests of linear relationships between opinions and causal priors generally impose the assumption that the observations are independent, and the random samples on which opinion surveys are administered ensure that this assumption is met. This widely used methodology imposes a "blind spot" for the possibility that respondents' opinions were influenced by their peers.[9] Without data on the opinions of peers, researchers have little choice but to search for "within-individual" explanations among other attributes of the respondent, as if each respondent arrives at an opinion independently of what others around them are thinking.[10]

[9] We are grateful to an *AJS* reviewer for suggesting the "blind spot" metaphor.

[10] We use the term "within-individual" to include the formation of shared collective identities that might involve historical processes unfolding in large populations. The process remains within-individual if the identity of each member of the collective can be attributed to the effects of some other attribute of that individual, including membership

Examples of studies that suffer from the atomistic blind spot are easy to find; indeed, the authors have not been immune, as shown by previous work attributing the liberal political views of the "new middle class" to incumbency in intellectual occupations (Macy 1988). We are in good company. Kohn's path-breaking work on occupational self-direction (Kohn 1969; Kohn and Schooler 1983) and Stouffer's (1955) research linking education with liberal tolerance of nonconformity are seminal studies that have shaped subsequent opinion research for the past half century. More recently, Inglehart (1997) explains changing social mores within and across populations by pointing to the within-individual effects of increasing education and affluence.

Rather than a consequence of demographic and moral organizing principles, a parsimonious relational alternative is to view alignment as the emergent property of a self-organizing cultural landscape. This approach not only addresses the puzzle of lifestyle politics but also helps to resolve long-standing controversies about the spatial and social clustering of political and lifestyle preferences (Lipset and Rokkan 1967; Bishop 2008). The relational explanation is surprisingly simple: the self-reinforcing dynamics of two well-documented social processes—homophily (McPherson, Smith-Lovin, and Cook 2001) and social influence (Marsden and Friedkin 1993; Christakis and Fowler 2007, 2008). Social influence refers to processes by which individual attributes are shaped by exposure to others (Mark 1998, 2003; Kitts 2006; Christakis and Fowler 2007, 2008; Flache and Macy 2011*a*, 2011*b*). Although influence is often equated with the tendency to become more similar, the dynamic can also be negative, as when in-group members seek to differentiate from members of an out-group (Sherif 1966; Tajfel and Turner 1986; Hogg, Turner, and Davidson 1990; Brewer 1991). In short, influence includes mechanisms that lead people to become more similar, through "diffusion, contagion, imitation, assimilation, cooptation, convergent competition or a host of other processes—or more dissimilar—through repulsion, divergent competition, differentiation, etc.—as a result of interaction" (Dow, Burton, and White 1982, p. 173).

Homophily refers to processes of selection (including self-selection) such that "ties of amity" (Kitts 2006) are more likely to form between similar persons (Doreian 1989; McPherson et al. 2001). As with influence, while homophily is often equated with the formation of positive social ties, the reverse is also possible, in which "ties of enmity" are more likely to form between those who are dissimilar (Kitts 2006). Homophily and influence become self-reinforcing when the attraction to those who are similar and

in the group. We distinguish "within-individual" from "atomized," by which we mean the assumption that individuals are not influenced by network neighbors. This distinction is important for the relational model we propose that incorporates both "within-individual" and "between-individual" influences.

differentiation from those who are dissimilar entail greater openness to influence. The result is network autocorrelation—the tendency for people to resemble their network neighbors.[11]

McPherson (2004) argues that homophily and social influence among demographically similar individuals can generate an "ecology of affiliation" in which political and cultural measures become differentiated in "Blau space." McPherson illustrates the relational explanation with the example of status differences in support for gun control: "The best explanation . . . is simply that individuals who are like one another communicate their likes and dislikes, their values and beliefs, their attitudes and opinions to one another. As a result of this process, these opinions often become localized in social space" (p. 273). This explanation has a provocative implication: Even an observed partial correlation between a demographic attribute and an opinion, controlling for a host of other demographic and individual attributes, may still be largely, if not entirely, spurious. McPherson continues: "The bottom line for this [relational] mode of explanation is a very radical one, from the point of view of traditional survey analysis: most relationships between variables in survey analysis are due to the fact that similar people talk to one another. All the elaborate 'causal models' of attitudes and opinions and such need to rule out this baseline explanation before they can convincingly argue that there is anything going on in their data beyond social transmission among similar others."

The relational explanation reverses the causal arrow linking geographic location to lifestyle suggested by previous studies (Marsden et al. 1982; Lamont et al. 1996; Gelman et al. 2008). While regional conditions may shape inhabitants' political and cultural preferences, individuals may also actively seek out locations in spatial networks that are amenable to their lifestyle (Bishop 2008). Gays and lesbians, for example, flock to tolerant and socially liberal urban centers such as New York, Chicago, and San Francisco (Florida 2002). Indeed, a vibrant literature in both sociology (e.g., Sassen 1991) and economics (e.g., Tiebout 1956) has explored the numerous features of geographic locales that attract individuals with different beliefs, orientations, and lifestyle preferences. In sum, spatial homophily can create a powerful feedback loop that amplifies individual-level regional effects.

[11] Network autocorrelation is the focus of recent research (Aral, Muchnik, and Sundararajan 2009) that highlights the methodological challenge in observational studies of social influence that may be confounded by homophily. For example, the formation of lifestyle enclaves might be due to the attraction of like-minded people rather than to the contagiousness of lifestyle preferences. While homophily alone is not sufficient to produce lifestyle correlations, homophily can amplify the confounding effects of peer influence by reducing the probability that people will be exposed to alternative points of view.

We are not the first to call attention to a relational feedback loop that can generate or amplify correlated opinions. McPherson (2004) sketched the outlines of the "ecology of affiliation," and Macy and Flache (2009) formalized his ideas as a computational model. More recently, DiMaggio and Garip (2011) showed how network autocorrelation can reinforce within-individual differences associated with innovation adoption, leading to social inequalities far greater than one would otherwise expect. We build on and extend these studies, using a computational model to show how lifestyle politics can emerge through a path-dependent historical process and how very small demographic and socioeconomic effects on opinion can be amplified many orders of magnitude.

A COMPUTATIONAL MODEL OF NETWORK AUTOCORRELATION

Network autocorrelation is the relational counterpart to spatial and temporal autocorrelation. Autocorrelation means that an observation is dependent on other observations, where this dependence increases with proximity in temporal, spatial, and network location. "Autocorrelation is the technical term which means, within the regression framework, that some variable, or the error term, is correlated with itself, either directly or indirectly over time, through space or across a network" (Dow et al. 1982, p. 170). Less technically, spatial autocorrelation refers to the tendency for individuals to resemble others in their geographic community, colorfully illustrated by shared preferences for "books, beer, bikes, and Birkenstocks" among liberal residents of Portland, Oregon (Cortright 2002). Temporal autocorrelation refers to the tendency for individuals to resemble themselves over time, as in longitudinal studies of behavior and attitudes. Finally, network autocorrelation refers to the tendency for individuals to resemble their network neighbors (Christakis and Fowler 2007, 2008), either because similarity affects the formation of social ties or because people come to more closely resemble their neighbors or both.

Statistical methods exist to correct the estimation of standard errors of covariates among time-dependent observations (Newey and West 1987). However, network dependence poses an additional problem that is not so easily addressed. When a random sample is taken from a population in which observations are network autocorrelated, the social dependence cannot be measured in the sample since the sample observations are independent. Statistical controls for demographic similarities also fail to address the problem since network autocorrelation can cause opinions to become spuriously aligned with locations in Blau space, as McPherson (2004) warns. Simply put, the intractable problem is the data (or lack thereof), not the statistical model.

In the absence of network data in the GSS, we adopted a generative (Epstein 2006) rather than inductive analytical strategy.[12] The goal is to identify mechanisms capable of producing generically large correlations between opinion dimensions even in the absence of equivalent substantive linkages, corresponding to the broad pattern of correlation observed among lifestyle and political items in the GSS. While this generative approach cannot rule out alternative explanations, it offers far greater parsimony than correlation-specific within-individual explanations.

Using an agent-based model, we simulated McPherson's (1983, 2004) "ecology of affiliation" among individuals who come to resemble those in the same demographic niche and to differ from outsiders through the relentless dynamics of homophily and influence. McPherson's original ecological model (1983) was used to examine membership recruitment and competition between voluntary organizations. Organizations recruit members through social ties shaped by homophily and distance in sociodemographic space (or "Blau space"). Later, McPherson (2004) extended this approach to cover a wider variety of traits that are transmissible through network ties, including opinions.

Extending this model, we investigate the conditions in which inflated correlations are able to emerge among the attributes of the members of a network autocorrelated population. Following Watts (1999), we use a "connected caveman" graph in which each agent has k immediate neighbors in the same cave and the network consists of $n/(k+1)$ caves, where n is the population size. The connected caveman graph provides a highly stylized network structure that nonetheless captures the widely observed small-world properties of high clustering and low average path length. High clustering implies that if A and B are tied to one another, they will also tend to share mutual ties with C, D, E, and so on. Low average path length implies that any two individuals in the network can reach one another through a relatively small number of network ties, as famously captured by the expression "six degrees of separation." The contribution of small-world network theory has been to show how these seemingly incompatible properties nonetheless coexist in a wide variety of empirical network contexts. We set $k = 99$, approximating the cognitive limit to the number of people with whom one can maintain a stable social relationship (Dunbar 1992).

A chosen percentage Φ of the edges (or network ties) are randomly rewired using the degree-preserving procedure introduced by Maslov and Sneppen (2002). By "degree-preserving," we mean that ties are deleted and replaced by new ones in a manner that does not change any agent's

[12] Although the GSS network module asks questions about respondents' network neighbors, the neighbors are not included in the survey.

number of network neighbors. Take a simple example with four network "nodes" A, B, C, and D. Further assume that A shares an "edge" with (i.e., is tied to) B while C shares an edge with D. To "rewire" these ties, the Maslov-Sneppen procedure would simply swap them such that A becomes connected to D and C to B, thereby preserving the original degree of each node. Edges are selected at random from the network and rewired using this procedure until Φ percentage of the edges have been swapped. We set $\Phi = 10\%$, which Watts (1999) found was sufficient to produce the characteristic small-world condition of high clustering and low characteristic path length.[13]

Each agent has five static attributes corresponding to fixed demographic traits and 20 dynamic attributes corresponding to changeable attitudes (i.e., political opinions and lifestyle preferences).[14] Static demographic attributes represent dimensions in the sociodemographic Blau space described by McPherson (2004), while traits on dynamic dimensions correspond to opinions, beliefs, attitudes, or behaviors that are transmissible through network ties. For simplicity, these 25 dimensions are binary (as gender, favor/oppose, or whether or not the agent engages in some activity).[15] All dimensions are randomized initially.

Each agent has a tie to its neighbors with a positive or negative weight based on the difference between the expected and observed Euclidean distance measured over 25 dimensions. The expected distance is that which we observe when positions on all 25 dimensions are random. Observed distances less than expected yield positive weights, and those greater than expected are negative, corresponding to the homophily assumption that agents are attracted to those who are similar and distance themselves from those who are different. Less formally, we mean that agents are drawn to one another when they share greater similarities than one would expect in a world of randomly assigned traits. The distance $d_{ij,t}$ and weight $w_{ij,t}$ between agents i and j at time t are calculated as

[13] Figure A1 in the appendix graphically depicts a network with five caves and 10% rewiring. In robustness checks, we found virtually identical results with as many as 30% of ties rewired, which is consistent with previous studies of the highly robust properties of small-world networks.

[14] Although there are hundreds of issues included in the GSS, we assume that only a small portion are sufficiently salient to alter network ties or motivate social influence. The number of dimensions relative to the population size strongly determines the coordination complexity and thus the computational load because it determines the likelihood that the incumbent of a particular opinion profile will find others with the same profile. This ability to find others with opinion profiles similar to one's own provides an "anchor" that prevents agents from changing their views.

[15] We also tested—and found qualitatively similar results—for nominal dimensions with more than two values and continuous metric dimensions.

$$d_{ij,t} = \sqrt{\sum\nolimits_{m\in S}(s_{mi} - s_{mj})^2 + \sum\nolimits_{m\in O}(o_{mi,t} - o_{mj,t})^2}, \tag{2}$$

and

$$w_{ij,t} = E(d) - d_{ij,t}, \tag{3}$$

where m is the dimension, S is the set of static attributes, O is the set of dynamic dimensions, and $E(d)$ is the expected distance at $t = 0$ when all traits are randomly distributed.

We modeled social influence as an urn model (Dandekar, Goel, and Lee 2013) with 20 urns, one for each dynamic dimension. Each neighbor of a given agent places its opinion in the agent's urn with a probability given by the absolute value of the weight of the tie to that neighbor divided by the sum over the absolute values of the weights to all other neighbors. Once the urn is filled, an opinion is randomly chosen and assigned to the focal agent. The probability that agent i adopts neighbor j's opinion can be expressed as

$$p_{j,t} = \frac{|\ w_{ij,t}\ |}{\sum\nolimits_{A(i,k)=1} |\ w_{ik,t}\ |}, \tag{4}$$

where $A(i, k) = 1$ when agents i and k are connected and $A(i, k) = 0$ otherwise. In other words, the urn is most likely to produce opinions that are currently held by neighbors with whom the focal agent shares a strong (positive or negative) tie.

The fact that ties with positive or negative weights are both represented in the urn raises an important point about negative social influence. Previous studies based on empirical research, theoretical inquiries, and computational modeling support the assumption that influence can be negative (Elias [1939] 1969; Heider 1946; Cartwright and Harary 1956; Sherif 1966; Schwartz and Ames 1977; Bourdieu 1984; Tajfel and Turner 1986; Hogg et al. 1990; Brewer 1991; Macy et al. 2003; Mark 2003; Jager and Amblard 2005; Kitts 2006; Baldassarri and Bearman 2007; Haider-Markel 2007; Takacs, Flache, and Mäs 2014). A recent laboratory study confirms that influence between highly dissimilar persons is negative but that negative influence is also generally weaker than positive (Takacs et al. 2014; see also Mummendey et al. 1982; Struch and Schwartz 1989; Brewer 1999).

Our formalization is thus designed to account for both the presence of negative influence and its relatively weaker effects compared to positive influence. If a ball matching the trait of a negatively weighted neighbor is placed in the agent's urn, we then replace that ball with one randomly chosen from one of the alternative traits. For example, suppose possible opinions on a given issue are red or blue and a focal agent has a negative tie to an out-group member whose opinion on that issue is red. Then this red ball would either be removed from the agent's urn or, with a probability of 10%, be replaced by a blue ball. The 10% probability means that negative influence is relatively weak.[16]

At each iteration an agent is randomly chosen and its weights and states are updated. We used an asynchronous updating rule such that agents are updated one at a time on the basis of the current state of the world. An alternative synchronous method updates all agents simultaneously on the basis of the state of the world that all agents experienced at the end of the last round of updating. Asynchronous updating is generally preferred because of possible artifacts that can result from synchronous updating (Huberman and Glance 1993). At each iteration, both the dynamic dimension and agent to be updated are chosen with uniform random probability. This process is repeated until the states converge to a stochastically stable solution. We say "stochastically stable" because the probabilistic urn model never reaches static equilibrium in the manner that a deterministic model would.

We tested for convergence on this stochastically stable solution by measuring changes in the "structural dissonance" (equivalent to "energy" in statistical mechanics) that is trapped in the network. Dissonance is a function of opinion differences between otherwise similar network neighbors and agreement between otherwise dissimilar neighbors, in keeping with theories of balance in social psychology (Heider 1946). More formally, we quantify the dissonance (δ) present in the network as

$$\delta = \sum_{i,j \in N} \sum_{o \in O} w_{ij}(2|o_i - o_j| - 1). \tag{5}$$

Convergence was said to obtain when two conditions were met: (1) a statistically significant drop in dissonance followed by (2) no further drop between two consecutive samples of 100 data points taken over 1 million iteration intervals. We dropped the very small number of realizations that failed to reach a stable solution within the time constraints imposed by the available computational resources. The model was implemented in Java and executed on the Amazon EC2 cluster. (The source code is available from the last author on request.)

[16] Robustness tests showed that almost any nonzero negativity is sufficient to obtain stable disagreement. Using a related model of homophily and influence, Mäs and Flache (2013) found that stable disagreement is possible without negative influence if opinion dimensions are continuous (rather than binary) and positive influence causes agents to take more extreme positions through the exchange of additional supporting arguments.

MODEL RESULTS

Consider a brief and highly stylized example with just four focal agents: John, Beth, Phil, and Laura. John and Laura are liberals while Beth and Phil are conservatives. Otherwise, the four individuals initially have diverse and crosscutting views with regard to many other opinion dimensions, such as consumer preferences. John and Laura share a homophilous attraction based on shared liberal ideology, just as Beth and Phil will be attracted through their shared conservative ideology. Since attraction facilitates social influence, Phil finds himself inclined over time to adopt other opinions held by Beth (and vice versa), and this in turn increases their mutual attraction, and so on. Nevertheless, Phil and Beth continue to differ on fixed attributes (e.g., gender) and may even continue to disagree on some issues because of influence from their other neighbors. A similar dynamic unfolds for John and Laura.

The computational model simply scales up these processes to much larger numbers of agents and dimensions. Figure 3 reports the mean pairwise correlation magnitude among dynamic dimensions for networks ranging in size from 500 to 5,000 agents. As with the GSS analysis, we take the average of the absolute values of the pairwise correlations since it is the

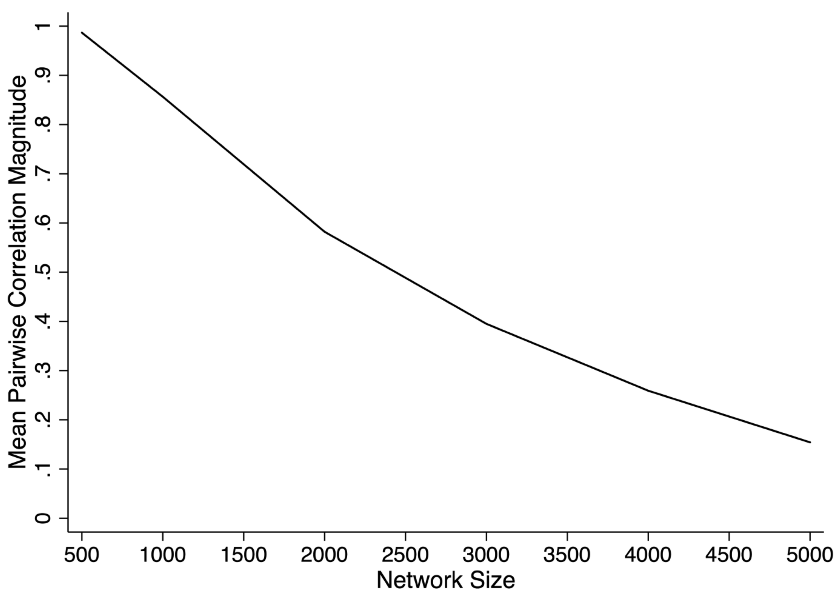

FIG. 3.—Weaker effects of network autocorrelation in larger populations: 20 dynamic and five static dimensions; 10% negativity; 10% rewiring; cluster size = 100. Values are plotted using a fractional polynomial function.

magnitude of the correlation that matters, not the direction. Because the model is stochastic, each condition was realized 100 times, and we report the central tendency based on a best-fitting fractional polynomial. (A time series illustrating a single realization is shown in fig. A2 in the appendix.) The results show that homophily and influence alone can generate generically large correlations, but only in relatively small networks. As network size increases, the correlations rapidly decline in magnitude. For the population size from which the GSS was sampled, figure 3 suggests that network autocorrelation alone may not be a plausible explanation for the level of empirically observed correlations.

As network size increases, low average correlations are the result of coordination complexity that increases exponentially with both the number of agents and their attributes. Given infinite time, opinions will eventually become aligned even with high coordination complexity—but perhaps not in our lifetime. In order for opinions to become correlated in finite time, coordination complexity imposes the need for a coordinating mechanism that aligns the population in a stochastically stable pattern. Two plausible candidates are (1) a within-individual effect of fixed attributes (e.g., demographic background) and (2) a between-individual effect of network "hubs" with very high in-degree (e.g., political and media elites). Figure 4 shows how a very small within-individual influence of static traits can coordinate the alignment of opinions in large populations. With probability p, the balls in an agent's urn are replaced by a ball corresponding to one of that agent's static traits. This generates an expected correlation between static and dynamic dimensions and between dynamic dimensions mutually influenced by the same static dimensions, plotted on the x-axis in figure 4.[17] The striking result is that the observed correlations on the y-axis are orders of magnitude greater than would be expected on the basis of their conditioning through p. The results show how even a very small within-individual effect is sufficient for inflated correlations to develop, regardless of network size. This superlinear mapping reflects the amplifying effects of homophily and influence, such that a small but consistent elective affinity between a sociodemographic trait and an opinion (e.g., social class and musical preference)—in the presence of homophily and influence—is sufficient to generate far greater issue alignment than if individuals arrived at

[17] The pairwise correlation between p and initial r (prior to social influence) is .892. Since static traits are orthogonal, the dynamic dimensions do not align equally across all static dimensions. For simplicity and interpretability, r is the largest expected correlation, corresponding to the static dimension that becomes most strongly aligned with the 20 dynamic dimensions. Although some demographic traits are correlated in natural settings (e.g., race and region), all static dimensions are orthogonal in the model (such as race and gender) as a more conservative assumption in a test of network autocorrelation.

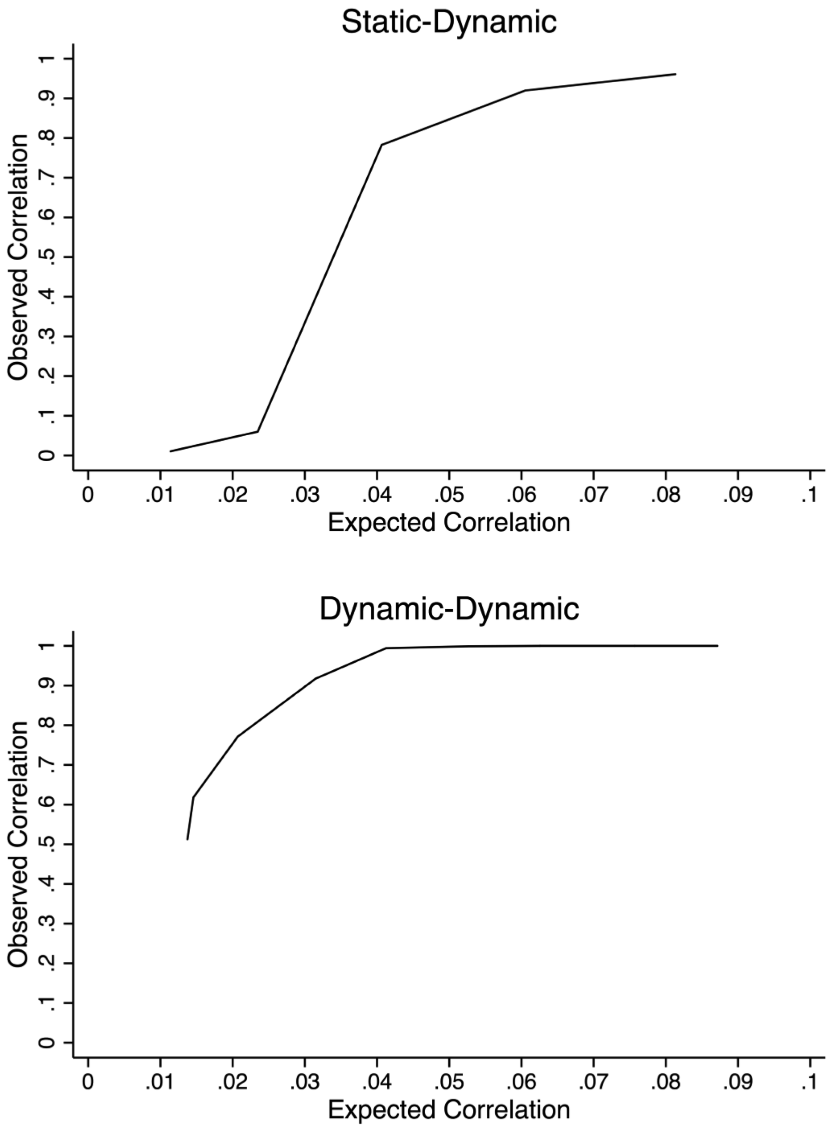

FIG. 4.—Amplified zero-order pairwise correlation by size of within-individual effect: 20 dynamic and five static dimensions; 10% negativity; 10% rewiring; cluster size = 100; N = 5,000. Values are plotted using a locally weighted scatter plot smoothing (lowess) function. The x-axis reflects the expected correlation between dimensions (in the absence of social influence) given the size of the within-individual effect of static dimensions on dynamic dimensions. Static dimensions condition dynamic dimensions through the probability p with which a neighbor's ball in the agent's urn is replaced with a ball corresponding to one of the agent's static traits. The y-axis measures the mean observed pairwise correlation magnitude.

their views independently of others. Moreover, opinions can become aligned even when some dynamic dimensions lack any substantive affinity for a static trait, as long as some of the others do.

The second coordinating mechanism corresponds to the top-down social influence of a small number of network hubs or "opinion leaders" (Lazarsfeld, Berelson, and Gaudet 1944; Watts and Dodds 2007). These network hubs could represent widely followed political leaders, radio or television talk show hosts, influential pundits, or political parties. Figure 5 shows how two opinion leaders to whom all agents are connected can facilitate the emergence of network autocorrelation, regardless of population size. The ties to these network hubs are unidirectional, with influence flowing only from the hubs to the rest of the population. However, the weights on these ties are updated in exactly the same way as the weights on all other ties. The static attributes and dynamic opinions of hubs are randomly distributed and remain fixed.[18] The strength of hub influence is implemented in the same way we varied the strength of within-individual influence. With probability p, an agent will update its opinion on the basis of hub influence rather than influence from neighboring agents. Figure 5 shows that—similarly to within-individual effects—a negligible amount of hub influence induces larger correlations than we would otherwise observe, given the coordination complexity of aligning multiple dynamic dimensions in large populations.

Note that the computational results often show correlations of far greater magnitude than those usually observed in the GSS. The reason is that our stylized model does not include idiosyncratic (or "noisy") influences on individual opinions. By incorporating noise, the model could produce much smaller correlations with closer correspondence to those seen in the GSS. However, the opposite does not hold: if the computational model produced correlations far smaller than those found in the GSS, then the model would not provide a plausible explanation for the empirical pattern.

DISCUSSION

So why do liberals drink lattes? The prevailing explanatory strategy in much opinion research looks for something about lattes that might bias hot-beverage choices within each independently choosing individual. For example, a consistent liberal preference for things foreign might plausibly explain the affinity for lattes, Volvos, and Beethoven, along with a class position that can accommodate a $5 cup of coffee and a $50,000 sedan. But

[18] As more than two hubs are introduced, average correlations decrease in magnitude because of crosscutting pressures from multiple influential sources who disagree with one another (see fig. A3 in the appendix). Figure 5 shows that the same is true when two competing hubs become increasingly strong in their influence.

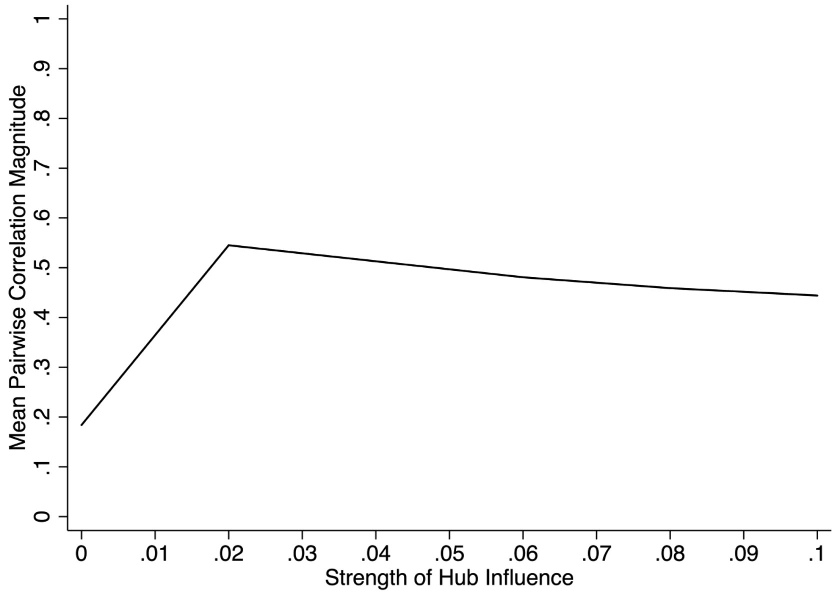

FIG. 5.—Coordinating effect of two opinion leaders: 20 dynamic and five static dimensions; 10% negativity; 10% rewiring; cluster size = 100; N = 5,000. Values on the x-axis are the relative influence of opinion leaders; with probability p on the x-axis, a neighbor's ball in the agent's urn is replaced with a ball from one of the hubs. Values on the y-axis are plotted using a fractional polynomial function.

how would an underlying cultural commitment to xenophilia explain the liberal predilection for home-grown mysticism and New Age spirituality?

Needless to say, our concern is not really about political differences in hot-beverage preferences. We provide this example to illustrate a more general conundrum for the social sciences. Data mining the GSS cumulative file reveals an important fact about social life as seen through the lens of survey research: correlation, not randomness, is the spontaneous condition of the social world.

It is well understood that these correlations can be causally spurious because of socioeconomic and demographic background. What is not so well understood is that the association between ideology and lifestyle preference can also be spuriously inflated because of network autocorrelation. Take, for example, the association between ideology and musical preference. For contemporary rock, gospel, heavy metal, and rap, more than 40% of the correlation with ideology evaporates when controlling for sociodemographic background. Crowther and Durkin (1982) found that musical preferences vary with age and gender, and these preferences can be plausibly attributed to developmental differences in emotional needs that operate within each individual, without regard to peer influence. Similarly, McCown et al. (1997) found a male preference for exaggerated bass that they attribute to gender differences in personality.

However, it is also possible that adolescents' musical preferences are shaped in part by the desire for social approval from gender- and age-specific friends (Selfhout et al. 2009), which then amplifies even very small developmental effects. More generally, just as correlated preferences can be spurious because of underlying demographic effects, so too can demographic correlations be vulnerable to spurious causation through the unmeasured effects of homophily and influence, such that ideology and lifestyle become differentiated in sociodemographic Blau space (McPherson 2004). Without data on network peers, opinion surveys offer limited ability to rule out network autocorrelation as a complementary or alternative explanation for demographic clustering that might otherwise be attributed entirely to within-individual effects of causally prior demographic attributes.

A long series of formal theoretical studies, dating back to Schelling's classic model of neighborhood segregation, have demonstrated that the emergence of culturally and ethnically differentiated clusters should not be assumed to reflect the aggregation of individual preferences and prejudices (Schelling 1971; Pancs and Vriend 2007; van de Rijt, Siegel, and Macy 2009). Our study was motivated by the possibility that a similar process of network autocorrelation may also explain the formation of lifestyle enclaves.

This possibility carries especially troubling implications when we can construct equally plausible explanations for a correlation and its opposite. Lazarsfeld (1949) famously lists six "obvious" findings from *The American Soldier* only to reveal that, in all six cases, the exact opposite was found to be true. Watts (2011, p. xiii) states the problem thusly: "when every answer *and its opposite* appears equally obvious, then, as Lazarsfeld put it, 'something is wrong with the entire argument of "obviousness."'"

Lazarsfeld's warning was confirmed in a recent online experiment. In the "music lab" study, participants randomly assigned to different groups were asked to rank songs downloaded from the Web (Salganik, Dodds, and Watts 2006). When participants could see the download counts for those in their own group but not the downloads for those in another group (who likewise could see only their own group's downloads), a consensus emerged in which the "winners" in one group might be the "losers" in another. For example, "Lockdown" by 52Metro ranked first in one group and near the bottom in another. The results challenge the "obvious" explanation that a strong consensus reflects an underlying reason for the similarity of preferences. Instead, the study showed that, when behavior is subject to peer influence, highly nonrandom patterns can be the path-dependent result of an initially idiosyncratic event. The results pointed to the highly counterintuitive possibility that picking the winners can actually be less predictable, the stronger the consensus. The reason is that social influence can lead to a stronger consensus compared to opinions that are arrived at independently, and social influence can also make the outcomes more path dependent and thus less predictable. Even a strong and persistent pattern might have turned out very differently if by chance the path-dependent process had taken a different turn at the outset.

Unfortunately, the GSS, like most survey data based on random samples, lacks the relational data needed to test for network autocorrelation as an alternative explanation for path-dependent self-organization in the alignment of opinions. We therefore used computer simulation to demonstrate how attitudes with vanishingly weak causal relations could become highly correlated. The model does not require the atomistic assumption that individuals arrive at their hot-beverage choices independently, pointing instead to the possibility that lifestyle preferences and political views become socially, spatially, and demographically clustered through the self-reinforcing dynamics of homophily and influence.

The absence of survey data on the opinions of respondents' neighbors also limits our ability to validate the computational model. Nevertheless, even in the absence of direct empirical confirmation, the model retains face plausibility by demonstrating the logical implications of two widely observed empirical phenomena. While there are few lawful regularities in the social sciences, homophily and influence would surely qualify as two of them (Mark 1998, 2003; McPherson et al. 2001; Christakis and Fowler 2007, 2008). The model's central prediction is also empirically confirmed, at least qualitatively, and might even qualify as an additional lawful regularity: a self-reinforcing dynamic driven by homophily and influence creates a cultural landscape marked by widespread correlation between opinion items, even when seemingly unrelated by content—not so different from what we discovered by data mining the GSS.

The modeling results point to a new understanding of the causal implications of static attributes like demographic background. In multivariate causal analysis, it is generally assumed that the observed effects of fixed attributes on opinion cannot be spurious, since a fixed attribute cannot be affected by an unmeasured prior (Davis 1985). That assumption is called into question by the spuriously inflated correlations our model generated between dynamic and static dimensions. The results suggest a new "blind spot," not unobserved heterogeneity but rather the unobserved network autocorrelation when opinions are studied using random samples that preclude the possibility of controlling for the opinions of network neighbors.

However, this is not to suggest that demographic or cultural backgrounds are unimportant. On the contrary, the modeling results point to a new understanding of the mediating mechanisms. When reverberated through the "echo chamber" of interaction with similar alters, even very small within-individual biases can serve as coordinating mechanisms that catalyze network autocorrelation in large populations. A similar coordinating role can be played by opinion leaders with broad influence, even if this influence is far weaker than that of peers. It takes only a very small "nudge," whether from "within" or "above," to tip a large population into a self-reinforcing dynamic that can carve deep cultural fissures into the demographic landscape. When cultural tastes in turn have a reciprocal effect on personal networks, such divisions are likely to be even further exaggerated (Lizardo 2006), leading to a starkly divided world of latte-sipping liberals and bird-hunting conservatives.

Although these divisions may have only a weak material foundation in shared experiences and interests, the consequences can be nonetheless real, not only for the people who get swept up in tragic and sometimes lethal conflicts but also for the social scientists who study them. When violent conflict erupts between culturally divided groups that differ demographically, it is tempting to assume that the groups have something "real" to fight over, that is, that the conflict is grounded in material interests, inequalities, or deep-rooted ethnic divisions. However, the computational model shows how these conflicts can also emerge through the self-reinforcing dynamics of homophily and influence with minimal substantive foundation. The tragedy of network autocorrelation is that even initially arbitrary cultural alignments can eventually generate opposing interests and identities that can give meaning even to a conflict that is largely an emergent property of a self-organizing process of political and cultural clustering. Simply put, while violent intergroup conflict is often attributed to strong ethnic identifications and prejudices, our model shows that it may be the other way around. Ethnic identities and rivalries may come to be highly salient among people who live in segregated cultural enclaves carved into Blau space by the relentless flows of influence and homophily.

By similar reasoning, historical explanations focusing on path-dependent processes of group formation and influence may be essential for understanding the processual mechanisms behind particular lifestyle correlations. For example, paisley dresses, long hair, and acid rock may have held meaning for youthful opponents of the Vietnam War precisely because these lifestyle preferences were different from—and disliked by—the "straight" world. Rather than viewing path-dependent outcomes as arbitrary, we see the construction of lifestyle preferences as culturally meaningful historical processes of group formation. Nonetheless, it is also highly plausible that the "Woodstock generation" might have shaved their heads if supporters of the war were the ones with long hair.

CONCLUSION: A "RELATIONAL MANIFESTO" REVISITED

Although the immediate aim of this study is to uncover an alternative to conventional explanations for the formation of lifestyle enclaves, there is a larger methodological agenda inherited from Emirbayer's (1997) "relational manifesto." Random sampling gives investigators the confidence that each survey respondent is independent of the others, which is necessary to obtain an unbiased representation of the distribution of individual traits in the underlying population. However, this independence also carries an important, if largely neglected, limitation: Unlike the members of the underlying population, the respondents in a national random sample are atomized individuals, unaccompanied by friends and family. In the absence of relational data, there is no way to measure the effects of sorting and influence in the clustering of opinions. Investigators are then left with only one analytical option: to assign all the explanatory power to other individual attributes (McPherson 2004). Kohn's (1969) classic study of "class structure and conformity" is a paradigmatic example: an individual's open-minded attitudes are the consequence of that same individual's education and occupational self-direction. "Members of different social classes, by virtue of enjoying (or suffering) different conditions of life, come to see the world differently—to develop different conceptions of social reality, different aspirations and hopes and fears, different 'conceptions of the desirable'" (Kohn 1969 p. 7).

A bias toward within-individual explanations may be methodologically inscribed in the use of random samples for most opinion research. Simply put, the tendency for opinion research to focus on within-individual explanations may reflect the inability to measure network autocorrelation in survey records that lack relational data. Furthermore, the generic tendency for opinions to be highly correlated not only with other opinions but also with demographic and cultural backgrounds means that the search for a within-individual explanation is almost certain to succeed.

The question is whether the explanation should be trusted. If opinions are highly vulnerable to network autocorrelation, then atomistic explanations that do not take into account the views of respondents' neighbors should invite the same skepticism as zero-order correlations without statistical controls for causal priors.

Admittedly, obtaining network data for large populations is notoriously difficult and expensive, but it is not impossible, as the Add Health project impressively demonstrates (Bearman, Jones, and Udry 1997). Moreover, we are now presented with an unprecedented opportunity to study opinion dynamics embedded in networks as a result of the increasing tendency for billions of people around the globe to interact using digital devices that record their interactions. Recent studies illustrate how data from social media can be used to measure the effects of homophily and influence on human behavior and interaction (Adamic and Glance 2005; Aral et al. 2009; Aral and Walker 2012; Bond et al. 2012). These early studies inspire us to close on a note of optimism, by pointing to the intriguing possibility that social science is poised for a new relational beginning.

Appendix A

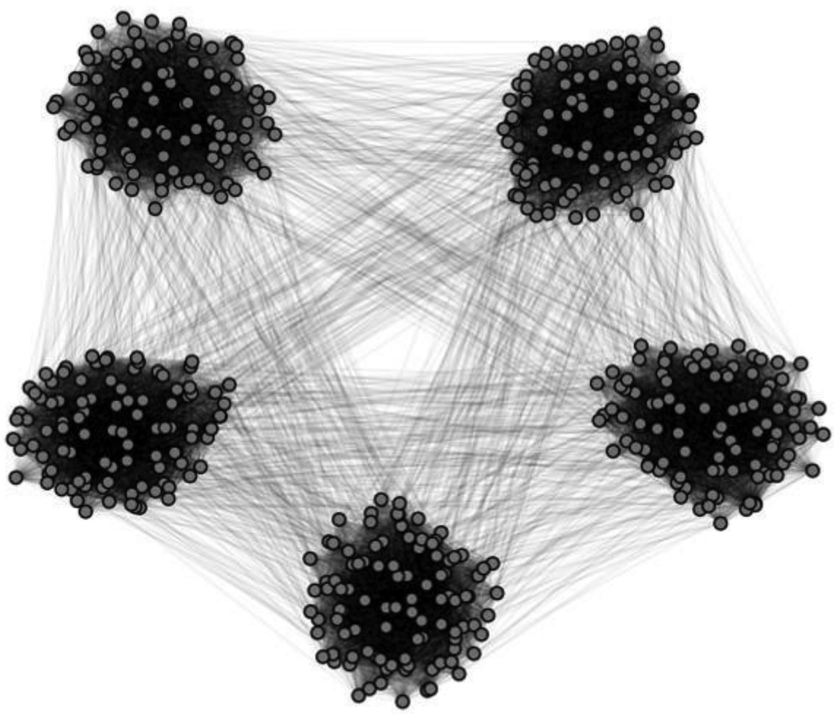

FIG. A1.—Graphic illustration of the connected caveman network. Depicted network features five caves with 100 agents per cave and 10% rewiring.

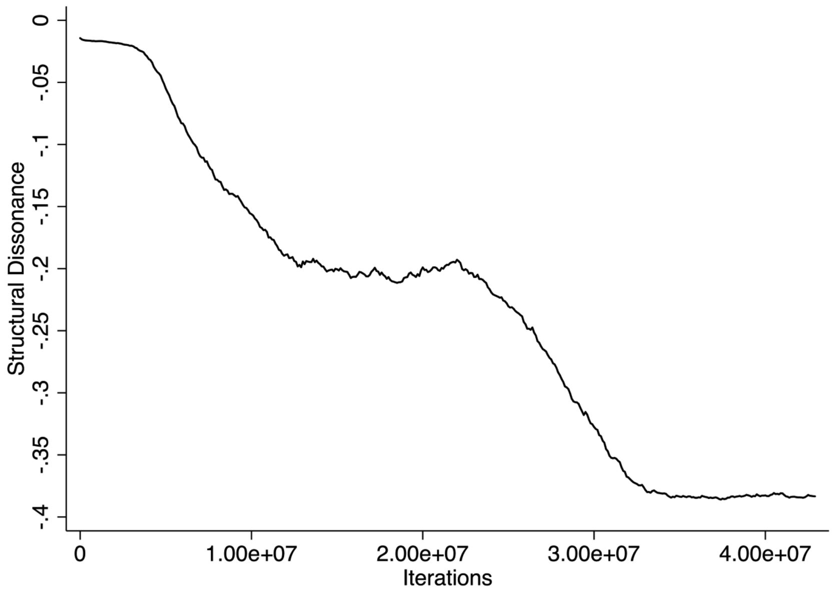

FIG. A2.—Representative time series of convergence to stochastically stable solution: 20 dynamic and five static dimensions; 10% negativity; 10% rewiring; cluster size = 50; N = 5,000.

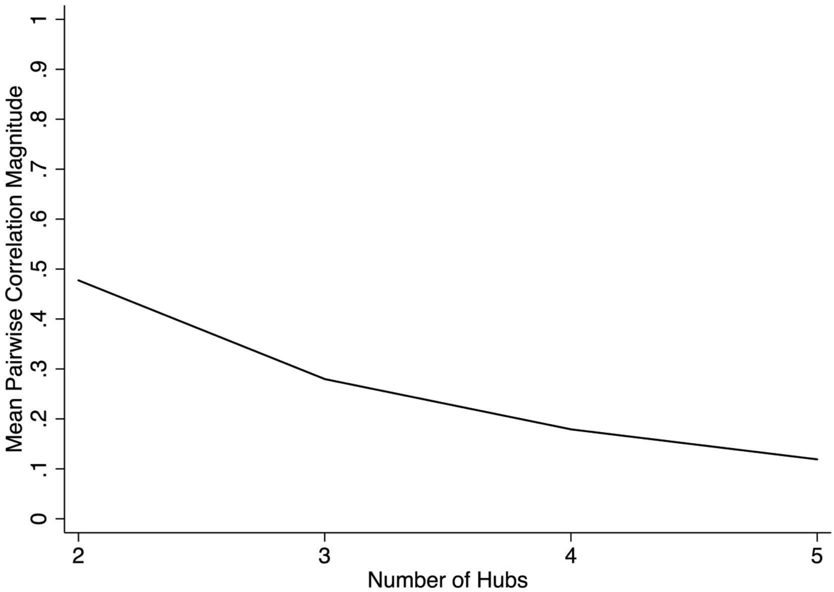

FIG. A3.—Decreasing correlation magnitude with increasing number of hubs: 20 dynamic and five static dimensions; 10% negativity; 10% rewiring; cluster size = 100; strength of hub influence = .05; N = 5,000. Values on the x-axis are the number of distinct hubs, from two to five. Values on the y-axis are plotted using a fractional polynomial function.

TABLE A1
MIXED-EFFECT ESTIMATES OF AVERAGE ZERO-ORDER AND PARTIAL CORRELATIONS BETWEEN PAIRS OF GSS LIFESTYLE ITEMS

	Zero-Order $\|r_0\|$	Partial $\|r\cdot\|$
Fixed effects:		
Intercept	.10***	.07***
Time (years)	−.00***	.00*
Residual SD:		
Time trends	.00	.00
Intercepts	.08	.06
Within pairs	.04	.04

NOTE.—N = 21,625 valid correlation-by-year observations nested within 10,180 lifestyle item pairs. Models are fitted using Markov chain Monte Carlo techniques. Time starts at 1972, and the intercept corresponds to the estimated mean correlation magnitude in that year. Estimates exclude 14 cases for which the reported correlation was based on three or fewer GSS respondents; the minimum N among the included cases is 171, with a median of 964 and a maximum of 2,956.

* .05 level.

** .01 level.

*** .001 level.

TABLE A2
MIXED-EFFECT ESTIMATES OF AVERAGE ZERO-ORDER AND PARTIAL CORRELATIONS BETWEEN 216 GSS LIFESTYLE ITEMS AND POLITICAL CORRELATES

	Political Ideology		Party ID	
	Zero-Order $\lvert r_0 \rvert$	Partial $\lvert r\cdot \rvert$	Zero-Order $\lvert r_0 \rvert$	Partial $\lvert r\cdot \rvert$
Fixed effects:				
Intercept	.07***	.05***	.02***	.01*
Time (years)	.00***	.00***	.00***	.00***
Residual SD:				
Time trends	.00	.00	.00	.00
Intercepts	.04	.03	.02	.01
Within pairs	.05	.05	.05	.05

NOTE.—N = 947 valid correlation-by-year observations for political ideology and 998 for party ID nested within 216 item pairs. Models are fitted using Markov chain Monte Carlo techniques. Time starts at 1972, and the intercept corresponds to the estimated mean correlation magnitude in that year. For party ID, we dropped respondents who answered "other party" since they cannot be placed on the Democrat-Republican scale.

* .05 level.

** .01 level.

*** .001 level.

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