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What Moves Partisanship?

Migration, State Partisan Environment Change, and Party Identification

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We consider how state political environments can alter the party identification and political behavior of individuals. Using panel data well-suited to assess the influence of migration on individual-level phenomena, we find that migrants alter their party identifications toward the majority party of their new states. Applying the estimates from this analysis to the 2000 presidential election suggests that individual-level change can alter presidential election outcomes in states if migration patterns meet certain conditions.

Keywords: *migration; party identification; political geography; state politics; political behavior*

The United States is a highly mobile society with 2.5% of the population moving from state to state annually (U.S. Census Bureau, 2006). As of 2000, only 60% of U.S. residents lived in their state of birth (U.S. Census Bureau, 2005). Given its prevalence, scholars have begun to assess the influence of migration on American politics. In particular, Gimpel and Schuknecht (2003) demonstrate the importance of migration for understanding presidential elections, showing that state-to-state migration affects how competitive parties are in states. For example, the only reason that the Democratic Party did not dominate presidential elections in California for the entire period since the 1960s is because migrants favored Republican candidates (Gimpel & Schuknecht, 2003, p. 84).

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Gimpel and Schuknecht (2003), however, do not address whether states' political environments affect the political behavior/attitudes of migrants. There is good reason to believe that state environments change migrants politically, though. The diversity of state politics is manifested in divergent cultures (Elazar, 1984), party systems (Mayhew, 1986), and variance in partisan and ideological composition (Erikson, Wright, & McIver, 1993). Moreover, political conditions within states affect the behavior of individuals. For example, state laws that subsidize electoral participation increase the likelihood that individuals vote (Francia & Herrnson, 2004; Wolfinger, Highton, & Mullin, 2005), and variance in states' systems for voter registration affects the likelihood that individuals identify with one of the major parties (Burden & Greene, 2000). In addition, the context of state politics conditions the degree to which voters' socioeconomic backgrounds (Jackson & Carsey, 1999a) and partisanship and ideology (Jackson & Carsey, 1999b) affect their votes in presidential elections.

Along these lines, we examine whether states' partisan environments affect the party identification of migrants. As such, our analysis follows Brown (1988) who shows that, at the county level, migrants alter their party identifications toward the majority party of their new environments. However, given the importance of states as political units, for example, in determining the outcome of U.S. presidential elections through the electoral college, it is essential to understand what role state environments have on the political behavior of individuals. At the individual level, party identification is one of the most influential factors in determining behavior, including individuals' votes (Bartels, 2000; Markus & Converse, 1979), their positions on issues (Carsey & Layman, 2006; Layman & Carsey, 2002), and their core values (Goren, 2005). Therefore, to the degree that state environments affect party identification, they have the potential to alter behavior in a manner that reinforces the prevailing viewpoints within them. In turn, this reinforcement at the individual level has the potential to affect political outcomes in the aggregate, such as which party's candidate wins statewide office and/or to which party's candidate a state's electoral votes are awarded.

Below, we develop a Bayesian perspective rooted in research on social communication to explain how migration may lead to change in party identification and assess whether such change occurs using data from the intergenerational panel study conducted by Jennings, Markus, Niemi, and Stoker (2005). Employing structural equation models to correct for measurement error in party identification, we find support for the explanation that state environments alter the party identification of in-migrants. In addition, we estimate the effect that changes in party identification have on individual vote choice and outcomes in presidential elections. This analysis suggests that although party identification is influenced by changes in state environments due to migration at the margins, this influence can be large enough to alter which party's candidate wins states' electoral votes.

Party Identification, State Political Environments, and Communication Networks

Party identification is conceived of as individuals' cognitive and affective orientations toward parties based on evaluations of objects associated with the parties (Campbell, Converse, Miller, & Stokes, 1960, Chap. 3). For example, the primary object driving individuals' orientations toward the Democratic and Republican Parties in the 1952 and 1956 surveys used by the authors of *The American Voter* was parties' associations with the Great Depression (Campbell et al., 1960, p. 45). Party identification was stable because voters' evaluations of the parties with respect to this event were stable, leading most voters to retain the same cognitive and affective evaluations of the parties over time. Subsequent research on the stability of party identification demonstrates similar stability (Green, Palmquist, & Schickler, 2002), reinforcing the view of party identification as an enduring orientation. Yet, party identification is not entirely stable. The authors of *The American Voter* themselves note that "personal forces," such as "a change in neighborhood" that exposes voters to situations in which the partisan leanings of their new neighbors are different from their previous neighbors, may foster change in identification (Campbell et al., 1960, p. 150), a hypothesis supported by Brown's (1988) county-level analysis.

Research on the effect of social communication on political attitudes provides a explanation for how migration can affect party identification in this way. In studying the influence of discussion networks on individuals' political attitudes, Huckfeldt, Johnson, and Sprague (2002, 2004) observe that individuals' main discussion partners exert greater influence on individuals' own attitudes when the discussion partners' attitudes are shared by a majority of the discussion network than when the main discussion partner is in the minority. One explanation for this influence is that when one confronts a new perspective that conflicts with one's preexisting attitude, one searches for new information to adjudicate the disagreement. When one does so within a discussion network in which most of the individuals favor the new attitude, most additional information will reinforce the new attitude, increasing the likelihood that one will replace one's preexisting attitude with the new perspective (Huckfeldt & Sprague, 1995).

The implication of this research for state-to-state migration is straightforward. By way of example, consider a Democrat from Massachusetts, a predominately Democratic state, who moves to predominately Republican Utah. Therefore, on average, a majority of individuals within the new network in which she becomes embedded will be more likely to hold, and express, views that are favorable to Republican, and unfavorable to Democratic, policies and candidates than if she remained in Massachusetts. Consistent with research on social communication, she will become more likely to perceive individuals in the network as Republicans (even if they are not; Huckfeldt, Beck, Dalton, Levine, & Morgan, 1998) and express attitudes consistent with her communication partners' (Republican) positions (Huckfeldt, Johnson, & Sprague, 2002, 2004).

Assuming that her new networks tilt in favor of Republican positions, the process through which partisan change occurs can be described using Bayes' rule. Consider that the Massachusetts transplant has an initial belief about the probability that she is a Democrat, Pr(D), that is driven by her cognitive and affective orientation toward the Democratic Party. On moving to Utah, the Massachusetts transplant may hear from a member of her communication network at her place of employment that Democrats favor giving people something for nothing—a reference to the Democratic Party's role in creating a welfare system that some Republicans argue created incentives among the unemployed not to work. In Bayesian terms, the transplant has a belief that Democrats favor giving people something for nothing given that she is a Democrat, Pr(S|D), and a belief that Democrats do so given that she is not a Democrat, Pr(S|~D). Using this information, then, the transplant updates her belief that she is a Democrat given the statement that Democrats favor something for nothing, Pr (D|S), in the following manner:

$$\Pr(D|S) = \frac{\Pr(S|D) \times \Pr(D)}{\Pr(S|D) \times \Pr(D) + \Pr(S|{\sim}D) \times \Pr({\sim}D)}.$$

What implication does this process have for party identification change? The answer is contingent on the conditional probabilities that the transplant attaches to Pr(S|D) and Pr(S|~D). If Pr(S|~D) > Pr(S|D), then the transplant's posterior, or updated belief, Pr(D|S), will be less than the probability that she attached to her identification with the Democratic Party, Pr(D), prior to hearing the statement. As a result, the information will make the transplant less certain that she is a Democrat. Critically, although the information that the transplant receives about the Democratic Party is negative, it is likely that Pr(S|~D) > Pr(S|D). To understand why, consider that the transplant is likely to assign a very low probability to Pr(S|D), because she is a Democrat. After all, her belief about the probability of negative information about the Democratic Party being true, given that she is a Democrat and holds a positive cognitive and affective orientation toward the Democratic Party, should be very low. However, the transplant also has some belief that she is not a Democrat, Pr(~D). That is, she may reason that she always thought of herself as a Democrat but that, in fact, she is not. If, in fact, she is not a Democrat, then, there should be a higher probability of negative information about the Democratic Party being true than if she is a Democrat. As a result, Pr(S|~D) should be greater than Pr(S|D) when she hears something negative about the Democratic Party, reducing Pr(D) over time.

Importantly, since the Massachusetts transplant will hear more negative information about the Democratic Party because of the move, over time, Pr(D) will decline even if the conditional probability, Pr(S|D), that Democrats favor giving people something for nothing is very low.[1] In addition, since the transplant will hear positive statements about the Republican Party more often than if she had not moved, she will gradually update an initial (low probability) belief that she is a Republican, Pr(R), in a manner that increases over time. In summary, the transplant is more likely to change her partisan identification, such that Pr(D) falls below .5 and Pr(R) increases above .5, than if she had not moved.

Assessing the Effect of Changes in State Partisan Environment on Party Identification

Green and Palmquist (1990, 1994) show that panel studies from the National Election Study (NES) used in past analyses to assess why individuals change their identifications overstate the degree to which change occurs (Brown, 1988; Fiorina, 1981; Franklin, 1984, 1992). After correcting for measurement error, Green and Palmquist emphasize that party identification is highly stable. Therefore, correcting for measurement error, we assess whether individuals from the youth component of the youth-parent socialization study (Jennings et al., 2005) change their party identification because of changes in their partisan environments. We focus on these respondents because Green and Palmquist (1994, p. 441) identify this group as the least stable panel that they analyze, which they attribute to greater pliability of partisanship among young adults (Krosnick & Alwin, 1989).[2] This relative instability makes this panel a good place to begin in examining whether changes in state environment affect party identification. That is, if changes in state environments do not affect individuals in this sample, such changes are unlikely to affect party identification generally.

We also focus on these data because respondents were interviewed over long periods with initial interviews in 1965 and subsequent interviews in 1973, 1982, and 1997. After all, in a panel study that samples individuals for more than a 2- to 4-year period, respondents who moved may not have lived in their current states for long enough to have their new environments affect their identifications. In addition, panel studies that interview respondents over short intervals will have a lower percentage of respondents who moved. The youth component of the intergenerational panel study, then, is particularly appropriate for our purposes.

Employing these data, we corrected for measurement error in the observed party identification variables by creating a latent party identification variable for each panel.[3] Since we had only one observed indicator of party identification in each panel, we set the variances of the error terms for the observed measures of party identification in each panel equal to one another and constrained the regression weight between observed party identification and latent party identification to 1. Furthermore, we assume that there is no correlation between the measurement errors across panels. These assumptions ensure identification of the models reported below (Wiley & Wiley, 1970) and are consistent with structural equation models of party identification reported by Green and Palmquist (1994) and Carsey and Layman (2006).

To measure how respondents' state partisan environments differ from one panel administration to the next, we employ presidential election returns. Specifically, we obtain the mean percentage of the two-party vote won by the Republican presidential candidates during elections proximate in time to the administration of the panels in the state where the respondent lived during a later panel and subtract from it the equivalent percentage from the preceding panel. For example, for the Massachusetts transplant discussed above who moved to Utah during the interim between 1973 and 1982, the state partisan environment change variable is constructed as follows: where UT refers to the percentage of the two-party vote won by Republican presidential candidates in Utah and MA refers to that won by Republican candidates in Massachusetts:

$$(UT_{1972} + UT_{1976} + UT_{1980} + UT_{1984})/4 – (MA_{1972} + MA_{1976} + MA_{1980} + MA_{1984})/4.$$[4]

Since Republican presidential candidates outperformed Democratic candidates in Utah compared to Massachusetts during this period, the observation is positive, indicating that the partisan environment in the new state of residence was more favorable for Republicans than in the old state.[5] We expect this variable to be positively and significantly related to the latent party identification variables in the 1973, 1982, and 1997 panels, because higher values on these latent variables indicate greater identification with the Republican Party.

In addition, it may be that simply moving to a new state with a different partisan environment is insufficient to alter partisanship. Rather, consistent with the gradual updating process described above, exposure to the new state environment over time may be necessary. Therefore, we interact the state partisan environment change variable with a variable measuring the number of years that individuals resided in the state in which they lived during the year of each survey.[6] We expect the coefficients for the interaction terms to be positive and significant because, as individuals live in their states for longer periods, they receive both a greater volume of negative communication regarding the party disadvantaged by the state's environment and positive communication regarding the party advantaged.[7] Table A1 in the appendix displays summary statistics for all of the observed variables employed in the analysis.

Before proceeding, we acknowledge that states are diverse politically in that there are predominately Democratic (Republican) areas in predominately Republican (Democratic) states. This fact limits the degree to which our measure of state partisan environment change captures the exposure to pro-Democratic/Republican information. This limitation makes it less likely that we will observe a relationship between state partisan environment change and party identification change, making the estimates reported below conservative from a hypothesis testing standpoint.

Migration and Party Identification Change

Table 1 presents findings from two models of party identification using latent measures of party identification to correct for measurement error. Model 1 in the first column presents estimates of the effects of prior party identification and state partisan environment change on party identification in 1973, 1982, and 1997. Model 2 in column 2 differs in that the effect of state partisan environment change on party identification is modeled as conditional on the number of years that respondents have lived in their states of residence. Although the chi-square statistics for the models presented in Models 1 and 2 indicate that the implied covariance matrix of the

Table 1
Maximum Likelihood Estimates of Measurement and Structural Models of the Effects Between Change in Political Environment and Party Identification

	1	2
Latent party ID → observed party ID		
1965	1.00 (.93)	1.00 (.92)
1973	1.00 (.92)	1.00 (.91)
1982	1.00 (.92)	1.00 (.92)
1997	1.00 (.94)	1.00 (.94)
Stabilities coefficients of latent party ID		
1965 → 1973	.41 (.45/.04)***	.41 (.45/.04)***
1973 → 1982	.73 (.69/.05)***	.74 (.70/.05)***
1982 → 1997	.81 (.71/.05)***	.82 (.72/.05)***
Predicted effects		
State partisan environment change (SPEC) ’65-’73 → 1973 party ID	.03 (.08/.01)**	
SPEC ’73-’82 → 1982 party ID	–.00 (–.01/.01)	–.05 (–.13/.02)
SPEC ’82-’97 → 1997 party ID	.03 (.05/.02)**	.00 (.00/.03)
Years in state 1982 → 1982 party ID		–.02 (–.03/.02)
Years in state 1997 → 1997 party ID		.01 (.02/.01)
SPEC ’73-’82 X years → 1982 party ID		.01 (.13/.00)**
SPEC ’82-’97 X years → 1997 party ID		.00 (.06/.00)*
Model fit statistics		
χ^2	80.44	189.31
NFI	.96	.95
IFI	.98	.97
RFI	.91	.91
TLI	.95	.94
RMSEA	.04	.05
N	981	981

Note: The format of the entries is as follows: unstandardized coefficient (standardized coefficient/standard error). To correct for measurement error in party identification, we use a single indicator measurement model to obtain the true values of party identification. We employ the standard assumptions proposed by Wiley and Wiley (1970) to fully identify the models. The conditional effects of political environment on party identification from 1965 to 1973 are not included in the analysis due to data limitations. The model fit statistics presented include the chi-square statistic (χ^2), the Bentler–Bonett normed fit index (NFI), Bollen’s incremental fit index (IFI), Bollen’s relative fit index (RFI), the Tucker–Lewis coefficient (TLI), and the root mean square error of approximation (RMSEA). The models were estimated using Amos 6.0.
*$p < .1$. **$p < .05$. ***$p < .001$, one-tailed tests.

models differs from sample variance-covariance matrix, the other goodness-of-fit statistics indicate that the models fit the data well. Importantly, the root mean squared errors of approximation (RMSEA) are less than the .05 cut-off in both models, indicating a good model fit (Schumacker & Lomax, 2004, p. 81). Table 1 presents a number of other goodness-of-fit statistics all of which indicate that the models fit the data well.

Following Carsey and Layman (2006), we refer to the coefficient estimates of the effect of prior party identification on current identification as stability coefficients. As in prior studies (Carsey & Layman, 2006; Goren, 2005; Green et al., 2002; Green & Palmquist, 1990, 1994; Layman & Carsey, 2002), the coefficients evince the stability of party identification. From 1965 to 1973, 1973 to 1982, and 1982 to 1997, the stability coefficients are positive and significant and are large in magnitude. However, the coefficients indicate a weaker link between prior and current party identification than observed in prior studies. The stability coefficient between 1965 and 1973 is .41, and the equivalent coefficients for 1973-1982 and 1982-1997 are .73 and .81, respectively. Substantively, the .410 coefficient predicts that strong Democrats and strong Republicans (a 7-point difference on the NES party identification measure) will be separated from one another by 2.87 points on the 7-point NES party identification scale in 1973, whereas a .90 stability coefficient (which is on the low end of stability coefficients typically observed in prior work) predicts a 6.3 point difference.

To some degree, the differences in findings relative to previous studies are expected, given Green and Palmquist's (1994) finding that respondents in this survey experienced lower levels of stability than respondents in any other panel study.[8] Nonetheless, these estimates suggest that party identification among the young is less stable than among the electorate as a whole and that it is less stable over longer intervals (8/9 to 15 years) than shorter intervals (2 years) normally examined in studies of party identification change. In fact, these findings are the first to examine party identification change during any one interval spanning more than a decade (1982-1997). As such, the lower stability coefficient (.810 for the 1982-1997 interval) for party identification is a notable finding in and of itself.

Shifting attention to the effect of migration, the positive and significant coefficients for the 1965-1973 state partisan environment change variable and the 1982-1997 state partisan environment variable suggest that individuals shift their partisanship with the prevailing partisan winds of the states to which they relocate. For example, several respondents moved to a state (actually the District of Columbia) where Democratic Party candidates faired on average 37.67 points higher in terms of the two-party vote than Republican Party candidates from 1980 to 2000 compared to their previous states. Similarly, one respondent moved to a state where Republican presidential candidates faired 17.29 points better than their Democratic opponents during this period (from Oregon to Idaho). Therefore, the .035 state partisan environment change coefficient predicts a 1.32 decrease (away from Republican identification) on the 7-point NES party identification scale for the individuals who lived in a partisan environment in 1997 that was so starkly pro-Democratic compared to that of their state of residence in 1982 (the 37.67 pro-Democratic difference). Similarly, the model predicts a .61 increase (toward Republican identification) on the 7-point scale when individuals relocated to a starkly more pro-Republican environment during this interval (the 17.29 pro-Republican difference). The magnitude of these changes due to migration are greater than changes in party identification due to maximum differences in individuals' attitudes on abortion, government services, and aid to African Americans observed by Carsey and Layman (2006) and family values and moral tolerance observed by Goren (2005). Although the typical effect of migration is not as sizeable, as most respondents do not move to states so starkly different from their prior states, the effect of a more typical relocation is comparable to the magnitude of the relationships observed by Carsey and Layman (2006) and Goren (2005). An individual living in a state 3.46% percentage points more favorable to Republican Party candidates in 1997 than their previous states in 1982 (a 1 *SD* increase in the state partisan environment change variable for the 1982-1997 interval) is predicted to experience a .12-point movement toward the Republican Party on the 7-point NES scale. As seen in the positive and significant coefficient for the 1965-1973 state partisan environment change variable, the findings for this period are similar. However, for the 1973-1982 interval, the state partisan environment change variable is not related to party identification in 1982 in the manner predicted. On the whole, then, the findings presented for Model 1 provide mixed evidence regarding the effect of migration on party identification even if the state partisan environment change variables for the 1965-1973 and 1982-1997 intervals suggest effects at least comparable in magnitude to the effects of issue positions and core political values in effecting partisan identification change.

Therefore, to explore the relationship between state partisan environment change and party identification in more detail, we interacted the state partisan environment change variables with the number of years that respondents lived in the states where they resided when the panel surveys were administered for the 1973-1982 and 1982-1997 intervals (data limitations preclude ascertaining how long respondents lived in the states in which they resided in 1973).[9] If the capacity of the partisan environment in which individuals live to change their party identification increases with the duration of their exposure to it, the coefficients of the interaction variable for state partisan environment change, and the number of years lived in the new state should be positively and significantly related to the latent party identification variables. In fact, as presented in column 2, we observe these relationships for the interaction variables. In addition, the substantive effects of state partisan environment change in the conditional model are similar to those discussed above. For example, the coefficient for an individual who lived in her state for 9 years in 1997 is .036, and the coefficient for an individual who lived in her state for the 15-year span is .06.[10] Therefore, Model 2 predicts that the migrant who moved from Oregon to Idaho would have roughly the same increase in the 7-point NES party identification scale (toward Republican) identification as the model presented in column 1 if that individual resided in Idaho for 9 years at the time of the survey. However, if she lived in Idaho for the entire 15 years between 1982 and 1997 after having lived in Oregon in 1982, the model predicts an even greater change in identification—an increase of 1.04 points in favor of Republican identification on the 7-point scale. Of course, the typical individual in 1982 did not reside in a state so starkly different in its support for presidential candidates compared to that of her previous state. A more typical effect during this period is given by considering the effect of a standard deviation increase of 3.42% points in the state partisan environment change variable between 1982 and 1997. An individual who lived in a state that supported Republican presidential candidates at 3.42% points higher than the state from which she moved 9 years earlier would be predicted by the model to move .12 points toward Republican identification on the 7-point NES scale; the equivalent figure for an individual experiencing 3.46% additional support for Republican candidates for 15 years is .21. The effects are similar for the 1973-1982 period for which the model provides a state partisan environment change coefficient of .040 for an individual living in their state for 9 years as of 1982.[11]

The Electoral Consequences of Migration Effects on Party Identification

Since migration leads to changes in party identification due to differences between the environments of migrants' new and old states, and as party identification affects voting behavior, we should observe changes in the voting behavior of migrants. The findings in the previous section, though, cannot speak to the magnitude of such changes or to their effect on electoral outcomes. Are such effects statistically significant, as we observe in Table 1, yet substantively meaningless in terms of their consequences for elections? Alternatively, do changes in party identification due to state-to-state migration reverberate from the individual to the aggregate level? We consider these questions by using our estimates of the effect of state partisan environment change on party identification for the 1982-1997 period for state-level presidential election outcomes in the 2000 presidential election. We focus on this period due to the closeness of the 2000 presidential election in many states. If individual change due to state-to-state migration can alter state electoral outcomes, then such changes should be observed in 2000.

We begin by estimating voting in presidential elections from 1980 to 2000, using data from NES studies for each election and a model employed by Nadeau and Lewis-Beck (2001) that is presented in Table A2 of the appendix. We then obtain predictions regarding for which candidate respondents are expected to vote based on the model. Obviously, the model includes a variable for party identification. Next, for each state decided by a margin of 3.5%, or less, in the two-party vote during the 2000 election, we reestimate the model 51 times (for the 50 states and the District of Columbia). In doing so, we alter the value of respondents' observations for the party identification variable, and the variables with which it is interacted, based on what Model 1 of Table 1 would predict for individuals moving to each state decided by 3.5% or less from every other state and the District of Columbia. For example, if respondents moved from New York to Florida, Model 1 of Table 1 expects a .352 increase, toward Republican identification, for the latent party identification variable.[12] This pro-Republican change occurs because Republican presidential candidates performed better in Florida than New York in the elections from 1980 to 2000. We then reestimated the model presented in Table A2 of the Appendix with each respondent's party identification observation increased by .352, obtaining predictions about the candidate for whom the New York-to-Florida model expects respondents to vote, allowing us to compare these predictions to the predictions from the original model in Table A2. This process allows us to arrive at a prediction of the percentage of voters changing their votes in the 2000 presidential election (in this case in favor of the Republicans) if they moved from New York to Florida—and from every other state (and the District of Columbia) to Florida. For example, comparing the predictions of New York-to-Florida model to the model presented in Table A2 leads to the prediction that 6.52% of voters moving from New York to Florida would change their votes from the Democratic to the Republican candidate based on the .352 shift in party identification toward the Republicans.

Table 2
Predicted Effects of Changes in Party Identification on State Vote Counts During the 2000 Presidential Election

State	Margin of Victory (in Votes)	Estimated No. of Migrants Changing Votes Toward Winning Party Due to Party ID Change	Change State Election Outcome?
States won by the Republicans			
Florida	537	27,097	Yes
New Hampshire	7,211	3,292	No
Nevada	21,597	5,168	No
Missouri	78,786	1,779	No
Ohio	166,735	3,383	No
States won by the Democrats			
New Mexico	366	–376	No
Iowa	4,144	610	No
Wisconsin	5,708	49	No
Oregon	6,765	1,504	No
Minnesota	58,607	3,234	No

Next, using data from the U.S. Census Bureau on the number of individuals moving from state to state from 1995 to 2000 (U.S. Census Bureau, 2003), we obtain a prediction for the number of individuals (6.52% of them) within the 1995-2000 migration cohort who changed their votes in favor of the (Republican) party winning Florida, discounting the prediction for the voter turnout rate in Florida.[13] This process created a count of the number of individuals estimated to change their votes in favor of the winning party because they moved from New York to Florida. We then repeated this process for all states and the District of Columbia and summed the estimates to obtain the number of voters predicted to change their votes in favor of the winning (Republican) party in Florida due to 1995-2000 migration. We present these predictions in Table 2, which itemizes the states decided by 3.5% of the vote or less by the winning party's candidate, the margin of victory in raw votes of the winning party's candidate, the number of votes that we estimate changed in favor of the winning party's candidate due to migration, and whether the outcome of the election changed in each state.

Beginning with Florida, the number of votes that we predict changed in favor of the Republican candidate, George W. Bush, due to relocation exceeds the number of votes by which he won the state. Therefore, our findings suggest that the effect of migration on party identification altered the candidate receiving Florida's electoral college votes and, obviously, in this election year, altered which candidate won the presidential election. Another way of stating this finding is to note that if changes in state partisan environment did not affect party identification then 27,097 voters in Florida who voted for Bush would have instead voted for the Democratic candidate, Al Gore, awarding the state and the election to him. In large part, this finding is driven by the closeness of the 2000 outcome in Florida. Given its closeness, one might argue, a change in virtually any factor affecting voting behavior could have altered the outcome. We find ourselves sympathetic to this critique. However, given the prediction of a change in 27,097 votes in favor of the winning party, we think it is clear that the effect of migration on party identification would have been great enough to alter an election decided by thousands more votes. This is especially the case as the predictions of changed votes are based on 1995-2000 migration totals only. Clearly, however, many more people moved to Florida during the period to which our 1982-1997 estimates can be applied than just those who moved between 1995 and 2000. As long as the migration patterns between the early 1980s and 1995 were largely consistent with the 1995-2000 patterns we used to produce these predictions, this fact makes these predictions more of a floor than a ceiling as far as the number of votes that were changed due to migration. Given these considerations, we feel comfortable with the finding that migration patterns to Florida, characterized by a large net of people relocating to Florida from states more Democratic than it, in combination with the effect of state partisan environment change on party identification, changed thousands of votes during the 2000 presidential election.

However, it is also telling that Florida is the only state where we predict the outcome was altered in the 2000 election. As noted above, this prediction is driven by a large net of relocation from states more Democratic than Florida. For example, the single largest migration from any one state to any other state during the 1995-2000 period was from New York (where from 1980 to 2000 the Democratic presidential candidate faired 11% points better on average than in Florida) to Florida when 308,230 people made this move (U.S. Census Bureau, 2003). In the other states where close races occurred, such a large volume of migration from dissimilar states did not occur. For example, consider the state decided by the fewest number of votes, New Mexico, where Al Gore prevailed even though his party lost a predicted 376 votes due to migration. From 1995 to 2000, the two states providing New Mexico with the largest number of migrants were California and Texas (U.S. Census Bureau, 2003). The effect of migration from California to New Mexico on voting favors the Republicans, because Democratic presidential candidates faired better in California than New Mexico from 1980 to 2000. However, migration from Texas, where Democratic candidates faired worse from 1980 to 2000 than in New Mexico, to New Mexico more than offset the Republican advantage that occurred due to California to New Mexico relocations. In other close states, such as Wisconsin, there was relatively little in-migration. In addition, most migrants to Wisconsin moved from bordering states, such as Michigan, where the state partisan environment was relatively similar to Wisconsin's. These factors, in combination, meant that migration could not change many votes in such states.[14]

Discussion

In summary, our findings with respect to the effect of the observed measures of state partisan environment change on the latent party identification variables indicate that states' partisan environments lead migrants to shift their party identifications in favor of the majority party, as operationalized by presidential election returns, in their new states. Although these effects do not lead to predicted changes in identification such that individuals alter their identifications from strong partisans of one party to strong partisans of the other, they are not expected to do so given that party identification is a stable phenomenon (Green & Palmquist, 1990, 1994; Green et al., 2002). At the same time, however, changes in state partisan environment exact noticeable changes in party identification that are similar to changes that occur due to policy positions (Carsey & Layman, 2006) and core political values (Goren, 2005). Given the importance of party identification for political attitudes and behavior, that such migration affects party identification means that it may affect electoral outcomes.

What is more, we observe precisely this effect in using our individual-level estimates of the effect of state partisan environment change on party identification from 1982 to 1997 to assess the effect of 1995-2000 migration on state-level outcomes in the 2000 presidential election. Although the factors necessary for changing large numbers of votes—a large volume of net migration favoring one party and a large number of migrants—are not present in most states, and although the number of states whose electoral college votes that we predict were awarded differently in the 2000 election due to the effect of state partisan environment change on party identification reduces to Florida, we feel that this finding is noteworthy. After all, in 2000, it only would have taken a changed outcome in one state to alter the outcome of the presidential election. As it happens, based on our analysis, such a change occurred in Florida, where state partisan environment change favored the Republicans due to aggregate migration patterns from predominately Democratic states to predominately Republican Florida, altering thousands of votes in an election decided by hundreds.

Given that elections have implications for the subsequent policies produced by government, we also think it important to recognize the importance of any factor that, potentially, can alter electoral outcomes. We will leave it to others to detail the policy consequences of the 2000 election. However, we believe it is realistic to think that many matters important to Americans, including outcomes involving war and peace and the nature of the tax code, would be different today if the Florida outcome had been different—an outcome we have argued was sensitive to the intermingling of migration, state level partisanship, and individual party identification. As such, we view these findings as being important to those wishing to build a comprehensive account of the political factors shaping important electoral and policy outcomes in the American democracy. This is especially the case due to the unit rule, which is employed by all but two states to allocate all of their electoral votes to the winner of a plurality of votes in their states. The rule guarantees that systematic changes in voting due to migration that lead to small margins of victory, as in Florida in 2000, can alter to whom an entire state's electoral votes are awarded, potentially altering the outcome of presidential elections.

In summary, this research stresses the importance of patterns of geographic support for the parties for understanding the political behavior of individuals and aggregate electoral outcomes. The perspective developed in this article attributes the effects of geography to the majority sentiment of citizens living within political units. Following Brown (1988), who examines county level migration, we focus on state-to-state migration due to the importance of states in the U.S. political system. Because new and different partisan environments affect party identification and behavior, it makes sense to believe that state-to-state migration has this effect. Our findings suggest that it does. In addition, these effects have the potential to influence aggregate political and, by extension, policy outcomes if the right conditions, such as migration patterns favoring one party over another and a sufficient volume of migration, are met.

Appendix

Table A1
Descriptive Statistics for the Observed Variables

Standard Variable	Minimum	Maximum	Mean	Standard Deviation
Party ID				
1965	0	6.00	2.52	1.84
1973	0	6.00	2.58	1.72
1982	0	6.00	2.81	1.81
1997	0	6.00	2.96	2.02
State partisan environment change				
1965-1973	–30.23	22.48	–0.01	3.91
1973-1982	–37.75	43.53	0.23	4.28
1982-1997	–34.07	16.00	0.19	3.18
Number of years lived in state				
1982	1.00	10.00	8.21	2.77
1997	1.00	18.00	14.31	3.90
Parent party ID	0	2.00	0.94	0.72
Race	0	1.00	0.07	0.25
Gender	0	1.00	0.49	0.50
Education	0	1.00	0.48	0.50
Income	0	23.00	12.69	6.82

Table A2
Logit Model of Voting for the Incumbent Party Candidate in Presidential Elections, 1980-2000

Variable	Estimate	Standard Error
National business index	0.021***	0.002
No incumbent running	2.647	1.792
National business index × No incumbent running	–0.149**	0.085
Incumbent party	–0.166	0.060
Party identification	0.036**	0.022
Party identification × Incumbent party	0.990***	0.022
Race	0.160	0.190
Race × Incumbent party	1.917***	0.190
Constant	–0.080	0.053
% correctly predicted	86.37	
Reduction of error	27.79	
Log-likelihood	–2465.84***	
Chi-square statistic	4674.66	
N	6,933	

(continued)

Table A2 (continued)

Note: The National Business Index (NBI) is constructed from responses to the Survey of Consumer Attitudes and Behavior during the last quarter of presidential election years, as presented by Nadeau and Lewis-Beck (2001, pp. 161-162). The NBI is the percentage of respondents indicating better in response to the following question minus the percentage indicating worse: Would you say that at the present time business conditions are better or worse than a year ago (Nadeau & Lewis-Beck, 2001, p. 161)? No incumbent running equals 1 if that is the case; 0 otherwise. Incumbent party equals +1 if the Democrats hold the White House; –1 if the Republicans hold it. Party identification is the 7-point National Election Study (NES) party identification scale rescaled from +3 (strong identifier of the incumbent party) to –3 (strong identifier of the out party) with 0 being those respondents indicating independence in not leaning on one of the parties. Race is coded 1 for non-Whites and 0 otherwise. The incumbent party coefficient, which represents the effect of the Democrats holding the White House for White independent voters, is significantly (p = .1; two-tailed) related to the probability of voting for the incumbent party and is negatively signed. However, although we have no hypothesis that such voters should oppose the Democrats under these circumstances—the p value for such a one-tailed test would be exactly .05—we do not note significance for this variable. Nadeau and Lewis-Beck also include a variable representing respondents' assessments of how well the economy is likely to perform in the near future; however, we do not include this variable, as it does not perform in the manner observed by the authors for the 1980-2000 period. We obtained these data from the NES cumulative file.
**$p < .05$. ***$p < .001$.

Notes

1. Of course, as Huckfeldt, Johnson, and Sprague (2002, 2004) note, individuals are embedded in multiple networks and the influence of one network can counterbalance the influence of another. It is not hard to imagine the transplant retaining ties to Massachusetts through family and friends that would reverse the decline in Pr(D) due to exposure to statements from the pro-Republican network. It is also not hard to imagine that the transplant becomes part of discussion networks in Utah that are pro-Democratic, as would no doubt occur in an employment-based network if the transplant moved to work on behalf of the Democratic Party of Utah. Nevertheless, the probability that the transplant becomes part of pro-Republican communication networks, and hears information critical of the Democratic Party, is higher due to the move than it would have been if she remained in Massachusetts.

2. We also replicate the analysis presented below in Table 1 for the adult portion of the panel study. Although the adult respondents in 1965 were not reinterviewed in 1997, the panel includes surveys in 1965, 1973, and 1982 only. The findings of the analysis for adults are similar in the manner state partisan environment change affects party identification. We do not present these findings to conserve space. The findings are available from the authors.

3. The observed party identification variables are the traditional 7-point identification scale employed by the National Election Study (NES), where 1 indicates strong Democrat, 2 indicates weak Democrat, 3 indicates independent-leaning Democrat, 4 indicates Independent, 5 indicates independent-leaning Republican, 6 indicates weak Republican, and 7 indicates strong Republican.

4. The election results employed for the 1965-1973 span were the 1964, 1968, 1972, and 1976 presidential elections; the results employed for the 1973-1982 span included the 1972, 1976, 1980, and 1984 elections; and the results employed for the 1982-1997 span were from

the 1980, 1984, 1988, 1992, 1996, and 2000 elections. By employing election returns for the same state across elections, we reduce bias from an election in which the candidate of one party did unusually well for that party in that state. By employing election returns across all states in a single year, we control for elections in which one party's candidate defeated the other party's candidate by a greater margin than it would normally, for example, 1964 and 1972. In other words, pooling the data in this way controls for idiosyncrasies in individual states in a given election and idiosyncrasies for all states in a given election. We include data from presidential election returns immediately before and after administrations of the surveys because such elections convey information about the state environment during the interval between administrations. For example, the presidential election in 1972 communicates information about the degree to which the partisan environment in states favored Republicans and Democrats during the interim between the 1973 and 1982 panels.

5. Scores of 0 indicate instances in which respondents lived in the same state during successive panel administrations. Theoretically, it is possible for an individual to move to a state that has an identical political environment to the state from which they moved, though this did not occur in our data.

6. We also include variables to control for respondents' gender, race, income, educational attainment, and the party identification of their parents. To conserve space, the estimates for these variables are not included in Table 1. These results are available from the authors on request. In the structural equation model, race is allowed to covary with educational attainment, income, and parent's party identification; educational attainment is also allowed to covary with gender and income. The base variables for partisan environment and number of years lived between administration of panels, and the interaction term for these variables, are also allowed to covary. Finally, the base variable for state partisan environment and the interaction term for state partisan environment and the number of years lived for one panel interval are allowed to covary with the variables measuring these phenomena in subsequent panels on the basis that many respondents live in the same state across multiple survey administrations; therefore, these variables are likely to covary across periods.

7. The models were estimated in Amos 6.0. Missing data is treated by employing full-information maximum likelihood estimation (FIML). The advantages of using FIML estimation over list-wise and pair-wise deletion are discussed by Arbuckle (1996). Briefly, and relying on the discussion of FIML in Enders and Bandalos (2001, pp. 433-435), FIML assumes that observations on the dependent variable are missing at random in that values of the independent variables affect the probability that the dependent variable is observed. FIML employs information about the correlation between the independent and dependent variable for observations for which both variables are observed to estimate values of the dependent variable when it is missing. Enders and Bandalos (2001) employ Monte Carlo simulations to show that FIML produces unbiased estimates that are more efficient than using list-wise deletion, pair-wise deletion, and similar response pattern imputation.

8. Green and Palmquist (1994) also examine the youth component of the intergenerational panel study (Jennings et al., 2005); however, they examine data from the 1982 follow-up panel, observing higher stability coefficients for the 1965-1973 interval (.577) and 1973-1982 (.951) periods. We attribute the differences in coefficients between their analysis and the coefficients presented in Table 1 to differences in the 1982 and 1997 samples.

9. Documentation from the 1973 youth–parent survey indicates that the residence of the respondent in 1973 should be the same as the last place the respondent mentioned when responding to questions about his/her state residences since 1965. However, of the 935 respondents that answered the series of questions that ask where the respondent has lived since the

last interview in 1965, 39.68% (371 observations) of the responses do not correspond with the variable that indicates the individual's residence in 1973. Therefore, we cannot be certain about the number of years each respondent lived in his/her state of residence in 1973, leading us to forgo an analysis using the interaction variable for this period.

10. Although the relationship between partisan environment and partisan identification is conditional on the number of years respondents resided in their states, there is a different coefficient for partisan identification across the range of values for the number of years respondents could have lived in their states (respondents could have lived in their states for between 1 and 15 years; therefore, there are 15 different coefficients for partisan environment). The base coefficient for partisan environment between 1982 to 1997, which provides an estimate of the effect of partisan environment when respondents lived in their state for 0 years (which is actually outside of the range of experience of the years lived variable), is .000, meaning that the model estimates that, for an individual who lived in their state for 0 years, the partisan environment of their state has no effect on their partisan identification. Therefore, for the 1982-1997 span, the conditional coefficient for partisan environment is (0) + (.004 × the number of years), where .004 is the coefficient for the interaction between partisan environment between 1982 and 1997 and the number of years lived in the state of residence in 1997.

11. The .040 coefficient is the conditional coefficient for an individual living in her state for 9 years, where the base coefficient for partisan environment is –.050 and the coefficient for the interaction of the number of years lived and partisan environment is .010: (–.050) + (.010) × (9).

12. The calculation is 11 × .032 = .352, where 11 is the result of subtracting the mean Republican presidential vote from 1980 to 2000 in New York (41.28) from that in Florida (52.28), and .032 is the coefficient for state partisan environment change from Model 1 in Table 1.

13. In using the estimate of the effect of state partisan environment change on party identification from Model 1 of Table 1 to arrive at a prediction of the number of votes changed in state level presidential elections due to migration, we make several assumptions. First, we assume that the coefficients from the 1980-2000 NES national samples of voters reflect the effects of the variables in the presidential vote choice model for voters in each state and the District of Columbia. Obviously, this introduces uncertainty into our predictions about the number of votes changed. However, it is the only option available to us because we must use the intergenerational panel study administered by the NES to assess the effect of state partisan environment change on party identification. Although we are using NES data, we employ the 7-point NES measure to construct the latent variable for party identification that serves as our dependent variable. Simply put, no state level voting data with a sufficient sample size exists containing this measure of party identification. Therefore, we must use the national sample from the 1980-2000 NES studies. Fortunately, in the one state in which our predictions indicate that the outcome was changed due to the effect of state partisan environment change on migrants' party identifications, the difference between the margin of victory for the winning party and the number of votes changed was so great that it is unlikely that this assumption affects the prediction. Second, we assume that states' turnout rates apply to all in-migrants similarly. That is, for example, migrants from New York and California (and every other state) to Florida are equally likely to vote.

14. To provide a sense of which voters are predicted to alter their votes, we also estimated the effects of migration on individual voting behavior across the range of our state partisan environment change variable. The results, available from the authors and not presented to conserve space, indicated that individuals responding that they were Independents in the 1997 panel were the voters who the models discussed above predicted would change their votes. (This is not surprising, as we did not expect any great apostasy of partisans due to relocation.) Therefore, the individuals predicted to change their votes in Table 2 represent Independents

moving from a state to another state with a different state partisan environment. Related to the finding that it is Independents who are predicted to change their identifications, one possibility that could confound the predictions presented in Table 2 is that individuals are more likely to move to states where the party with which they identify is the majority party. If, for example, New York Republicans are more likely to move to Florida than New York Independents, then there will be less party identification change due to migration. In turn, fewer voters will change their voting decisions. Having noted this possibility, there is no evidence that partisanship drives relocation decisions; in fact, research on state-to-state migration finds that economic factors affect internal migration (e.g., Davies, Greenwood, & Li, 2001). Although assessing the determinants of migration decisions is outside the scope of this article, we leave it to future research to determine whether migration is endogenous to the congruence between individual party identification and state partisan environment.

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