EXHIBIT 14
I. Introduction

In this report, I respond to issues raised by rebuttal declarations of Drs. Daniel Smith, Barry Burden, and Paul Gronke. Dr. Smith’s rebuttal focuses on my criticisms of the ecological inferences he drew from individual voting data aggregated to the census block. But as I discuss below, his responses often misrepresent my arguments and do not resolve the underlying issues. Therefore, I continue to have substantial concerns about his conclusions regarding the gap between early-in-person (EIP) voting rates of blacks and whites in Ohio.

Drs. Burden and Gronke focus on my claim that the scholarly literature on election administration does not support the plaintiff’s proposition that the availability of EIP voting raises aggregate or African-American voting turnout. Dr. Burden argues that his own work showing that EIP does not boost turnout does not apply to cases such as Ohio where EIP is reduced. However, no evidence is presented to support this argument. Second, Dr. Burden argues that his work implies that same-day registration manifest in Ohio’s “golden week” should
mitigate any negative effects of EIP voting. As I discuss below, there are reasons to question this claim.

Dr. Gronke, in particular, argues that much of the work I cited about EIP voting is outdated and superseded by studies of elections from 2008 to 2012. Below I address two concerns about Dr. Gronke’s report. First, he often overstates the differences between the more recent literature and the post-2008 studies. Both literatures support the argument that EIP voters are high-propensity voters and are thus likely to participate despite marginal changes in the early voting window. Second, none of the recent studies he cites provides compelling evidence that EIP boosts overall or African-American turnout. Nor do any of the studies show that reducing the EIP window will disproportionately reduce African-American voting.

Finally, having received access to the data underlying Dr. Smith’s report only after the filing of my initial report, I will also take the opportunity to address some issues surrounding the quality of his geo-coded voting data and the appropriateness of his ecological inferences.

II. Dr. Smith’s Response to My Criticisms of His Ecological Inferences

In his rejoinder, Dr. Smith responds to several of my criticisms of his attempt to draw inferences about individual EIP voting rates from aggregate data. Unfortunately, many of his responses mischaracterize my arguments and objections. For example, Dr. Smith quotes me at length “affirming” his regression results (Smith, page 3). In each instance, I am simply describing his bivariate relationships between the EIP rate and black VAP at the census block level. In no case, did I “affirm” his interpretation of what those correlations mean.
Below I respond on each of the points raised about ecological inference in his rejoinder.

A. Low Correlations

As I explained in my initial report, my concern about the low correlation between black VAP and the EIP voting rate is significant because the conditions under which one can reliably draw inferences about differential EIP voting rates from correlations are quite stringent. The conditions essentially require that there are no omitted variables correlated with race or the racial composition of a block that also explain EIP voting rates. That there is such a weak correlation between EIP rates and black VAP increases the likelihood that there are many omitted variables such as related demographics and partisan strategies that explain variation in the usage of EIP voting opportunities. They are many reasons to believe that these variables may also be correlated with race and the local racial context.¹

B. Omitting Racial Mixed Neighborhoods

In his response, Dr. Smith downplays my criticism of the omission of racially heterogeneous census blocks by asserting that Ohio has such significant residential segregation that there are few voters in such blocks.

Using the 2012 census block data provided by Dr. Smith, I can compute the number of EIP voters who live in census blocks that are not 0% or 100% black (i.e. those excluded from the homogeneous block analysis). I calculate that such blocks account for 336,339 EIP voters, more

¹ While Dr. Smith does report the correlations, the extent of the low correlations was visually obscured in his scatter plots by his decision to truncate the plots at EIP voting rates of .25. Based on his 2012 data files, this truncation omits approximately 20% of the EIP votes cast in 2012. EIP rates ranging as high as 1.0 can be found in both homogeneously white census blocks and homogeneously black census block.
than 60% of the total statewide. Using a looser definition of racially heterogeneity based on those in blocks ranging from 10% black to 90% black, I compute that such blocks have 142,440 EIP voters or about 26% of the statewide total. It seems imprudent to base conclusions on analyses that exclude such a large portion of the available data. In analyses below, I demonstrate the potential pitfalls of excluding the racial mixed blocks.

C. County Level Analysis

Dr. Smith badly misrepresents my intentions by including the scatter plot of EIP rates against black VAP at the county level. I did so precisely to illustrate the problem of aggregation bias, not to argue that whites utilize EIP voting at rates greater than blacks. At issue is whether one can conclude from Dr. Smith’s correlation analysis that blacks utilize EIP voting at rates greater than whites. The conditions for which such an inference is valid are the cases where there would be none of the aggregation bias that Dr. Smith describes. But it is precisely under those conditions that I should be able to aggregate the census block data to the county level and recover similar correlations. That I cannot suggests the presence of the sort of confounders that create the aggregation bias that undermine inferences from census block data.

I am far from the only scholar with concerns about aggregation bias and the reliability of ecological data to estimate early voting rates. A recent working paper by Ashok et al (2014), compares estimates of early voting rates by race using both individual data and census block data. Their concerns about the use of census block group data are worth quoting at length.

The early voting states for which we have individual-level race data reveal a substantially different pattern than the states for which we utilize block-group level data. In Florida, Georgia, and North Carolina, non-whites are much more likely to use early voting than
white voters, while the reverse pattern is implied by the ecological data. This may indicate that early voting turnout patterns by race are quite different in the South versus the non-South. However, it certainly indicates that we should be very cautious in making individual-level inferences from the Census block-group race data. Consequently, we focus less on race throughout the remainder of the analysis. (Page 17)

I am indeed aware that relationships can flip when data are aggregated. That is precisely my concern about the correlation analyses conducted by Dr. Smith at the census block level.

D. Analysis of Bounds

Dr. Smith concedes that his application of the method of bounds in his initial report was flawed. He now provides new figures based on the correct formula that takes into account the uncertainty association with the racial gap in overall turnout. He says that the conclusions of his earlier report are unaffected. That is not quite the case. In his previous report, the upper bound of the white EIP rate is generally around .1. Using the correction, the upper bounds are now around .15 across levels of racial homogeneity and the gap between the white upper bound and the black lower bound has shrunk. (Smith rebuttal, Figure 4a).

In Figure 5a of his rebuttal, Dr. Smith reports the adjusted analysis for EIP voting rates on the eliminated days. Strikingly, the bound for the white and black rates overlap for every level of homogeneity less than .96. Dr. Smith does not comment on the discrepancy between the corrected results and those of his initial report.
E. Probative Value of Five Counties in 2010

In his initial report and rebuttal reports, Dr. Smith argues that his analysis of the EIP voting in Ohio’s five largest counties has “probative value.” Alternatively, I argued that such analysis had little value as those counties are not representative of the state of Ohio as a whole. Unfortunately, we do not have the ability to look at differences between the five largest counties and the smaller counties in 2010. But we can examine the difference between the two sets of counties in terms of their 2012 behavior. Table 1 illustrates some key differences between the largest 5 counties and the remaining 79 counties in Dr. Smith’s 2012 census block data set.

<table>
<thead>
<tr>
<th>Table 1: Comparison of Five Largest Counties with Smaller Counties in 2012</th>
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<tbody>
<tr>
<td></td>
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<tr>
<td>----------------------</td>
</tr>
<tr>
<td>Black VAP</td>
</tr>
<tr>
<td>White VAP</td>
</tr>
<tr>
<td>EIP Rate</td>
</tr>
<tr>
<td>EIP Rate (eliminated days)</td>
</tr>
</tbody>
</table>

The first two columns report the turnout-weighted averages of the census blocks in each set of counties. The final column reports the differences. All of these differences are significant with p-values less than .001.²

The upshot of Table 1 is that Dr. Smith’s 2010 analysis likely excluded a set of counties for which the both the white VAP and EIP voting rates were statistically significantly higher than those counties included in his analysis. Thus, the top 5 counties represented a biased sample and therefore conclusions about the racial differential in EIP voting from 2010 are not reliable.

² These p-values do not account for clustering (see the discussion below). If I account for geographic cluster at the census tract level, the p-values for EIP rate, black VAP and white continue to be less than .001. The p-value for the difference in EIP rates during excluded days is .106 when clustering at the tract level is accounted for.
III. New Concerns about Dr. Smith’s Ecological Inferences

In this section I address several new concerns about Dr. Smith ecological inferences based on my analysis of the 2012 census block data that was provided.

A. Clustered Observations

In his reports, Dr. Smith reports that his estimates of the relationship between EIP voting rates and Black VAP are “statistically significant.” But I believe that the level of statistical confidence in his estimates may be inflated.

In a standard regression analysis, the researcher asks the computer to estimate the standard error of the coefficient on the dependent variable. This is a measure of the uncertainty surrounding the estimated value of the coefficient. Without further instruction, the computer will assume that all of the observations in the data set are statistically independent and the magnitude the standard errors will diminish at a rate proportional to the inverse of the square root of the sample size. For Dr. Smith’s application, the number of census blocks is well in excess of 200,000. So if one assumes that all of the observations are independent of one another, the standard errors will be very small and statistical significance is almost assured independently of the magnitude of the relationship between the EIP voting rate and the black VAP.

But in geographic data such as those based on census blocks, the assumption of the independence of observations is simply untenable. Outcomes from census blocks are almost assuredly correlated with outcomes in neighboring census blocks as well as with census blocks within the same municipal jurisdictions. For example, because election administration in Ohio is
carried out at the county level, there are many reasons to believe that outcomes across census blocks within the same county should be highly correlated.

The problem is that if we ignore such “clustering” the estimates of the standard errors will be artificially low and the likelihood of obtaining statistically significant “false positive” relationship increases. Dr. Smith does not appear to take account of these issues.

To demonstrate these problems, I examine the effects of clustering on a specific measure of statistical confidence in a regression coefficient, the 95% confidence interval. The confidence interval is simply the interval between the estimated coefficient minus 1.96 times its standard error and the coefficient plus 1.96 times the standard error. The 95% confidence interval of an estimated regression coefficient is the set of values for the true coefficient that we cannot reject at the conventional significance level of 5%. Thus, if the confidence interval is larger, there is a larger set of statistically plausible values for the true relationship.

In Table 2, I report the confidence intervals for the coefficient of EIP voting rates on black VAP for two of Dr. Smith’s regression models. Under Dr. Smith’s assumptions about the validity of ecological inferences, these coefficients reflect the racial differential in EIP voting. Therefore, the confidence intervals reflect the statistically plausible values for the racial differential under Dr. Smith’s maintained assumptions.

The first model that I examine is his regression of EIP voting rates and black VAP for the entire early voting period of 2012. The first row shows the confidence intervals for the model where geographic clustering is ignored. It implies that the true regression coefficient of black VAP on the EIP rate lies between .121 and .125 with 95% confidence. Alternatively, it implies that we could reject any racial differential in the EIP voting rate of less than .121. I also estimate
the model based on EIP voting during the days that would be eliminated for 2014 (row 4).

Without adjusting for clustering, the plausible range of estimates of the racial differential in EIP voting rates lies between .025 and .027.

But the situation changes substantially when one accounts for the possibility of geographic clustering. Based on the data provided by Dr. Smith, I am able to account for clustering at two levels: the census tract and the county. I believe clustering at the county is more appropriate given that elections are administered at that level. But I have reestimated the confidence intervals adjusting for clustering at both the county and census tract level. These are reported in rows 2 and 3 for the model of all EIP voting and rows 5 and 6 for EIP voting on the eliminated days. Note that the lower bound of the confidence interval decreases markedly especially when the clustering at the county level is accounted for. In the case of overall EIP voting rates, a racial differential as low as .057, or less than half of that estimated without clustering, cannot be rejected by the data. For EIP voting on eliminated days, a differential of .006 (e.g. 6 EIP votes per thousand votes), or less than one quarter of that estimated without clustering, cannot be rejected.
Table 2: Confidence Intervals with Different Assumptions About Clustering

<table>
<thead>
<tr>
<th>Model</th>
<th>Lower Bound of Confidence Interval</th>
<th>Upper Bound of Confidence Interval</th>
</tr>
</thead>
<tbody>
<tr>
<td>All Days (no clustering)</td>
<td>.121</td>
<td>.125</td>
</tr>
<tr>
<td>All Days (cluster at tract level)</td>
<td>.112</td>
<td>.134</td>
</tr>
<tr>
<td>All Days (cluster at county level)</td>
<td>.057</td>
<td>.188</td>
</tr>
<tr>
<td>Eliminated Days (no clustering)</td>
<td>.025</td>
<td>.027</td>
</tr>
<tr>
<td>Eliminated Days (cluster at tract level)</td>
<td>.022</td>
<td>.029</td>
</tr>
<tr>
<td>Eliminated Days (cluster at county level)</td>
<td>.006</td>
<td>.045</td>
</tr>
</tbody>
</table>

Clearly, however, even after accounting for clustering, the data do reject the hypothesis of no relationship between EIP voting rates and black VAP (zero is never in the confidence intervals). But my analysis shows that once clustering is accounted for, the data are not inconsistent with a very small effect -- even if Dr. Smith’s assumptions about valid ecological inference were true.

B. Non-Linearities: Another Diagnostic for Ecological Inferences from Bivariate Relationships

With the data provided by Dr. Smith, I am able to conduct an additional diagnostic test of the assumptions that underlie his ecological inferences from bivariate regressions on census blocks.

An important assumption required for appropriate inferences about individual behavior from bivariate aggregate regression models is that the relationship between the two variables be
linear (Gelman et al 2001). In the context of Dr. Smith’s report, this assumption requires that the average increase in EIP voting rate associated with a 0.1 increase of the black VAP from 0.2 to 0.3 should be the same as that associated with a 0.1 increase from 0.6 to 0.7 (or any other such increase).

A simple example illustrates why such an assumption is important. Suppose the true EIP rates for whites and blacks were 0.1 and 0.2, respectively. Then if the aggregate relationship between EIP rates and black VAP were driven by the differential utilization rates of the different races, each 0.1 increment in black VAP would be associated with the same 0.01 increase in the EIP rate. This is because each time we increase black VAP by such an increment the EIP rate should change only because there are now 0.1 more black voters who utilize EIP voting at rate 0.1 higher than the white voters they are hypothetically replacing. So the effect should be 0.1 x 0.01 or 0.01. If we did not find such a linear relationship, it would indicate that the aggregate relationship between EIP voting rates and black VAP was driven either by confounding factors other than race or that the racial differential is different for different types of census blocks. In such a case, we would not be able to use the bivariate regression model to estimate the differential EIP usage rate for the state as a whole.

There are many ways to test for linear relationship in bivariate regression models. The two most common are higher-order polynomial models and “piece-wise” linear models. Both approaches generate similar results for my analysis, but I will present the results of the piece-wise linear models because they are somewhat easier to understand and interpret.

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3 These are approximately the rates found by Dr. Smith in homogeneous blocks.
A piece-wise linear model is one that allows the linear relation between two variables to differ across different values of the dependent variable (in this case black VAP). In the models that I estimate, I allow the linear relationship to differ across each 0.1 interval of black VAP. In other words, the model estimates separate regression coefficients for census blocks with black VAP between 0 and 0.1, for census blocks with black VAP between 0.1 and 0.2, etc. If the assumption of linearity held, these coefficients would be all the same (or at least approximately). The appendix contains a table of the piecewise linear models that I estimate. This table contains both the regression coefficient for each interval and the estimated standard error after accounting for clustering at the county.\footnote{I get qualitatively similar results clustering at the census tract level and when clustering is ignored.} One can see that the regression coefficients do vary substantially across different levels of black VAP.

Figure 1a shows the predicted relationship between 2012 EIP voting and the black VAP from my piecewise linear model. The middle black line represents the expected EIP voting rate for each level of black VAP and the surrounding grey lines represent the 95% confidence intervals around those estimates. The figure provides strong evidence against the assumption of linearity. While there is a positive relationship between the EIP rate and the black VAP for low levels of black VAP, there is no statistically significant relationship between the two variables beyond a black VAP of 0.5.\footnote{Informally, this can be verified by noting that the expected EIP voting rate for a 100% black VAP block lies within the confidence interval of the rate at 50% black VAP. For a formal test, I estimated the simple regression model using only the census blocks with greater than 50% black VAP. The estimated coefficient is a very small .017 which is insignificant if observations are clustered at the county or tract level and has a two-tailed p-value of .068 if observations are not clustered at all.} Inconsistent with the required assumptions for valid ecological inferences, my results imply that the black VAP could be doubled without changing the EIP voting rate. It is not only the case that the positive relationship disappears beyond a black VAP
of 0.5, the relationship is negative over the black VAP segments from 0.8 to 0.9 and from 0.9 to 1.0.

**Figure 1a**

![EIP Voting Rate and Black VAP Piece-Wise Linear Model](image)

In Figure 2a, I present a similar analysis on the EIP voting days that are to be eliminated in 2014. The results are very similar to those of Figure 1a. There is a positive relationship between the EIP voting rate in the eliminated week and black VAP at low level of black VAPs, but none for high levels. At the very highest levels where black VAP exceeds 0.8, the relationship is negative.

My analysis of the non-linearities in the relationship between the EIP voting rate and black VAP is important for two reasons. First, it shows that the assumptions for drawing valid inferences about individual behavior are likely to be invalid. Second, the analysis demonstrates
that the racial patterns of EIP voting must logically be different in census blocks that are racially heterogeneous. This fact calls into question any attempt to extrapolate the results from homogeneous block analysis and the method of bounds to the state of Ohio as a whole.

**Figure 2a**

![EIP Voting Rate and Black VAP](image)

**IV. The Quality of the EIP Voting Data**

My examination of the 2012 Census block data used in Dr. Smith’s report has unearthed several data quality issues.

1. Presumably due to errors in geo-coding and/or the voter files that there are a number of Census block for which the number of EIP votes exceeds the total number of votes. Since these census blocks are estimated to have an impossible EIP voting rate greater
than 1, they were discarded for purposes of his report.\textsuperscript{6} My calculations reveal that there were 1606 such blocks and 5046 EIP voters that fell into this category.

2. The 2012 data reveal other obvious geo-coding problems that are not accounted for in Dr. Smith’s analysis. In addition to the one identified above, these include:

- The number of EIP votes exceeding the number of registered voters
- The number of EIP votes exceeding the voting age population.
- Total turnout exceeding the voting age population.

Then number of Census blocks with one or more of these problems is 19,012 and accounts for 35,220 EIP voters. This number represents over 6\% of the EIP voters state Drs. Smith and Herron successfully geo-coded in for 2012.

3. Over 4000 EIP absentee votes appear in Dr. Smith’s data file on days in which there was no EIP voting during 2012. Table 2 lists these dates and the number of EIP votes coded as occurring on each.

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\textsuperscript{6} I have verified this by examination of the R programming files that were provided. This fact is acknowledged in passing in Smith footnote 19.
### Table 3: EIP Votes on Non-Early Voting Days

<table>
<thead>
<tr>
<th>Date</th>
<th>Number of EIP Votes in Smith Data File</th>
</tr>
</thead>
<tbody>
<tr>
<td>Saturday, October 06, 2012</td>
<td>106</td>
</tr>
<tr>
<td>Sunday, October 07, 2012</td>
<td>9</td>
</tr>
<tr>
<td>Monday, October 08, 2012</td>
<td>186</td>
</tr>
<tr>
<td>Saturday, October 13, 2012</td>
<td>1439</td>
</tr>
<tr>
<td>Sunday, October 14, 2012</td>
<td>1009</td>
</tr>
<tr>
<td>Saturday, October 20, 2012</td>
<td>742</td>
</tr>
<tr>
<td>Sunday, October 21, 2012</td>
<td>5</td>
</tr>
<tr>
<td>Saturday, October 27, 2012</td>
<td>598</td>
</tr>
<tr>
<td>Monday, October 28, 2013</td>
<td>56</td>
</tr>
<tr>
<td><strong>Total</strong></td>
<td><strong>4150</strong></td>
</tr>
</tbody>
</table>

4. The errors described above are only those obvious enough to be detected by simple inspection of the data provided by Dr. Smith. The complete set of errors may be significantly larger.

These problems come on top of the fact that Drs. Smith and Herron were only able to successfully match about 91% of the EIP absentee votes to census blocks. Thus, around 9% of the EIP votes cast are not included in Dr. Smith’s analysis and another 6% appear to have been miscoded. The precise consequences of these problems for Dr. Smith’s conclusions are difficult to discern. Also missing from Dr. Smith’s analysis are the EIP vote cast in the four counties for
which no EIP data was available. In the 2012 election these counties accounted for 1.59% of the overall turnout.

V. The Literature on EIP Voting and Turnout

A common refrain in the rebuttal reports is that I and other defendants’ experts had relied too heavily on an older, outdated literature on early voting. Although most of the pieces I cited were published in the past 10 years, they argue that the world of early voting was upended around 2008 and the older findings do not apply.

There are at least two problems with these arguments. To the extent to which the plaintiffs’ experts want to focus only on the last six years raises a concern that the new findings are driven by unusual electoral contexts such as the first African-American presidential candidate or partisan strategies to mobilize around early voting rather than what we should expect over the longer term. There may also be novelty effects that wear off over time (Giammo and Brox 2010). The role of early voting in partisan mobilization strategies may also change in the future as has been the case for so many campaign tactics of the past. In other words, we have little way of knowing whether or not patterns of EIP usage will revert to the older patterns over time.

A second and more important concern is that the plaintiffs’ experts overstate the differences between the older literature and papers based on more recent elections -- many of which are not published or peer-reviewed. For example, Dr. Gronke now claims that the typical user of EIP voting is no longer a high propensity voter. Yet in a paper he cites on the 2008 election (Alvarez et al, 2012), the authors say:
Specifically, the probability of early voting is greater among liberal, well-educated, older, male, strong partisan voters. This fits much of the theory on convenience voters, in the sense that older people, strong partisans, and the well-educated are typically “likely voters.” [quotation marks in original] p. 256.

In his recent conference paper with Charles Stewart on the early voting in Florida from 2008 to 2012 (Gronke and Stewart 2013), Dr. Gronke writes:

Among early voters, there is very little day-to-day correspondence in voting between 2008 and 2012. The one exception is people who voted on the first day of early voting in 2008 (a Monday) who disproportionately voted on the first day of early voting in 2012 (a Saturday). This probably reflects the fact that early voters tend to be stronger partisans and more interested in politics than Election Day voters. [italics mine]

Moreover, Gronke and Stewart (2013) find that early voters in one election are the least likely to fail to vote in subsequent elections even in the case of Florida where the EIP voting periods was shrunk.7

On page 13 of his rebuttal report, Dr. Gronke says

I know of no empirical argument by which one could conclude that voters will, as a general matter, successfully adjust to the elimination of early voting and same day registration opportunities.

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7 Table 2 of Gronke and Stewart shows that early Florida voters from 2006 and 2010 respectively were the least likely to fail to vote in 2008 and 2012 respectively. In fact, despite the changes to early voting in Florida, early 2010 voters were more likely to vote in 2012 than were 2006 early voters in 2008.
But his own work with Stewart does provide some evidence that many voters did adjust to changes in the availability in the duration of early voting in Florida. From their abstract:

We demonstrate that while shortening the time of early voting does not appear to have hindered the earliest of early voters, it does appear to have dissuaded turnout among the latest early voters, especially those who previously voted on the final Sunday before Election Day.

Given that early voting Sunday before election day has been restored in Ohio, the results of Gronke and Stewart are consistent with the argument that Ohio’s changes should not have large effects on the use of EIP voting.

In his report, Dr. Burden cites two studies as evidence that restricting the time period of EIP voting disproportionately reduced African-American participation. The first is the Herron and Smith (2014) study which as I explain in my initial report does not actually show that changes in Florida’s EIP voting rules reduced African American turnout (although it did reduce African-American utilization of EIP voting).

The second paper cited by Dr. Burden is an unpublished paper by Glyn and Kashin (2014). But this paper is not principally about EIP voting. It is a paper proposing a new statistical methodology for estimating the causal effect of certain types of policy innovations. EIP voting in Florida is one of the examples the authors use to demonstrate the methodology. The gist of the EIP application is the demonstration that those voters who used EIP voting in one election are more likely to utilize it in the future and that this persistence is higher than other modes of voting. Under some implausible assumptions that the authors do not even attempt to
defend, these findings might be interpreted to suggest that EIP voting has a causal effect on turnout.\(^8\)

Notably in his rebuttal, Dr. Smith does not address the fact that his own work on Florida shows that:

- Most voters who cast early votes on eliminated days in Florida returned to cast ballots in the subsequent election.
- Blacks who voted on eliminated days were more likely than whites to vote in the subsequent election.
- The changes to early voting in Florida appear to have had no discernible effect on the racial composition of the electorate as the black percentage of the electorate was slightly larger in 2012.\(^9\)

VI. Same Day Registration

In my initial report, I did not directly address the literature on election-day registration (EDR) or same-day registration (SDR). As the plaintiffs’ experts argue, the literature does tend to show a positive effect of SDR on overall participation and Burden’s work shows that does mitigate his estimates of a negative effect of EIP. But Ohio is an extremely unusual case of

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\(^8\) The key assumption is that voters who utilized EIP voting in the earlier election have similar voting propensities as those who vote via other methods. As I have discussed, the literature finding that EIP voters are high propensity voters casts considerable doubt on this assumption. The authors also add the following caveat: “the estimates presented in this application are confined only to those individuals that utilized EIP in a previous election and hence we cannot comment on the overall turnout effect.”

\(^9\) Dr. Gronke does address this last point, but slightly misrepresents my argument. He says I argue that the changes in EIP had no affect on black turnout. I did not refer to black turnout but the black percentage of the electorate. My observation is simply that no evidence is provided to suggest Florida’s had a disproportionate effect on African-American electoral participation.
SDR. Ohio’s “golden week” was created by the concurrence of a thirty day registration deadline and 35 days of early voting. Therefore, same day registration ended thirty days before the election. Burden et al list 11 states that combined SDR and EIP in 2008. I list them here with their implied deadlines for SDR in order of those deadlines.\textsuperscript{10} As one see, Ohio is quite an outlier.\textsuperscript{11} In every state except New Mexico, a voter can register and vote within 15 days of the election. In most cases, this can be accomplished within a week of the election. Ohio and New Mexico are apples to these oranges. Therefore, any generalization about the conjunction of SDR and EIP from Burden et al may well not apply to Ohio.

VII. Conclusions

In my initial report, I highlighted several concerns with Dr. Smith’s attempts to estimate the racial differential in EIP voting rates for Ohio in 2012. I continue to believe my initial concerns were amply justified. My analysis of the census block level data from 2012 raises several new concerns about the validity of Dr. Smith’s inferences and the quality of the underlying data.

Second, I do not believe that the literature on early voting based on more recent elections is any more helpful to the plaintiffs’ arguments that the literature that I cited in my original report. The literature still lacks compelling evidence that EIP voting raises overall turnout or that

\textsuperscript{10} The registration deadlines were obtained and cross-checked from on-line sources such as \url{www.votesmart.org} and \url{http://www.mytimetovote.com/} as well as websites of state secretaries of state. I did not compile data on the early voting period for each state, so the underlying assumption is that registration deadline represents to most proximate time to the election that same-day registration can occur.

\textsuperscript{11} New Mexico was listed in Figure 2 as one that allowed early voting and same day registration. But in New Mexico, early voting begins the day registration ends. There is grace period up to the Friday after registration ends in which registration forms may be handed in at the County Clerk’s office.
of African-Americans. Studies of the effects of restricting early voting also do not draw clear conclusions about its effects on African-American turnout.

Finally, despite the arguments of the plaintiffs’ experts, I think it would be unwarranted to draw any conclusions about Ohio’s elimination of same day registration due to the vast dissimilarity between Ohio’s “golden week” and conjunction of SDR and EIP voting in the other states.

<table>
<thead>
<tr>
<th>State</th>
<th>End of EIP/SDR Period</th>
</tr>
</thead>
<tbody>
<tr>
<td>Ohio (Golden Week)</td>
<td>30 days prior to election</td>
</tr>
<tr>
<td>New Mexico</td>
<td>28 days prior to election</td>
</tr>
<tr>
<td>California</td>
<td>15 days prior to election</td>
</tr>
<tr>
<td>Iowa</td>
<td>10 days prior</td>
</tr>
<tr>
<td>Vermont</td>
<td>6 days prior (Wednesday preceding election)</td>
</tr>
<tr>
<td>North Carolina</td>
<td>3 days prior to election</td>
</tr>
<tr>
<td>Illinois</td>
<td>Election Day</td>
</tr>
<tr>
<td>Maine</td>
<td>Election Day</td>
</tr>
<tr>
<td>Montana</td>
<td>Election Day</td>
</tr>
<tr>
<td>Wisconsin</td>
<td>Election Day</td>
</tr>
<tr>
<td>Wyoming</td>
<td>Election Day</td>
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</tbody>
</table>
Bibliography


Herron, Michael C. and Daniel A. Smith. 2014 “Race, Party, and the Consequences of Restricting Early Voting in Florida in the 2012 General Election.” *Political Research Quarterly* Online first:

[http://prq.sagepub.com/content/early/2014/02/21/1065912914524831](http://prq.sagepub.com/content/early/2014/02/21/1065912914524831)
Appendix

Table A1 provides the estimated regression coefficients and standard errors for the piece-wise regression models discussed in Section III.B. The estimated standard errors have taken clustering at the county level into account.

<table>
<thead>
<tr>
<th>Coefficient</th>
<th>EIP Rate</th>
<th>EIP Rate (Excluded Days)</th>
</tr>
</thead>
<tbody>
<tr>
<td>BVAP (0 to .1)</td>
<td>0.195</td>
<td>0.039</td>
</tr>
<tr>
<td>BVAP (.1 to .2)</td>
<td>0.102</td>
<td>0.043</td>
</tr>
<tr>
<td>BVAP (.2 to .3)</td>
<td>0.139</td>
<td>0.025</td>
</tr>
<tr>
<td>BVAP (.3 to .4)</td>
<td>0.207</td>
<td>0.046</td>
</tr>
<tr>
<td>BVAP (.4 to .5)</td>
<td>0.168</td>
<td>0.009</td>
</tr>
<tr>
<td>BVAP (.5 to .6)</td>
<td>0.162</td>
<td>0.087</td>
</tr>
<tr>
<td>BVAP (.6 to .7)</td>
<td>0.063</td>
<td>-0.023</td>
</tr>
<tr>
<td>BVAP (.7 to .8)</td>
<td>0.160</td>
<td>0.028</td>
</tr>
<tr>
<td>BVAP (.8 to .9)</td>
<td>-0.147</td>
<td>-0.019</td>
</tr>
<tr>
<td>BVAP (.9 to 1)</td>
<td>-0.124</td>
<td>-0.069</td>
</tr>
<tr>
<td>Constant</td>
<td>0.097</td>
<td>0.012</td>
</tr>
<tr>
<td>R-Squared</td>
<td>0.066</td>
<td>0.032</td>
</tr>
</tbody>
</table>

Standard errors clustered by county in parentheses
I declare under penalty of perjury under the laws of the United States that the forgoing is true and correct to the best of my knowledge.

Dated: August 7, 2014

Nolan McCarty, PhD