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Does Income Inequality enter into an Aggregate Model of Voter Turnout? Evidence from Canada and Indian States*

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Abstract

Conflict theory argues that greater income inequality induces greater political and electoral participation. Relative power theory argues that greater inequality leads to political alienation and electoral disengagement. We test these alternatives on time series data by entering the Gini coefficient into an aggregate model of electoral participation in Canada. While ordinary least squares (OLS) results suggest that income inequality is inversely related to voter turnout, time series considerations raise the possibility that this result is spurious. Correction using a linear autoregressive distributed lag (ARDL) model finds no evidence of a relationship, but nonparametric modeling suggests an inverted U-shaped shape that is captured quadratically within the ARDL model. Additional support is found when the nonlinearity hypothesis is tested on a panel of Indian states. Together the results are consistent with the hypothesis that conflict theory operates at low levels of income inequality before growing inequality leads to voter alienation and lower turnouts consistent with relative power theory.

Key words: voter turnout, income inequality, Canadian time series and ARDL modeling, Indian state panel data, nonlinear relationships.

JEL: D72, D78, H62

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1. Introduction

In this paper we examine empirically whether a rise in income inequality, as measured by an increase in the Gini coefficient, increases or decreases voter turnout in Canada and in Indian state elections. This question, posed more generally as a test between the conflict and relative power theories of turnout, has been widely addressed in the literature but with outcomes that have remained empirically ambiguous (Stockemer and Scruggs, 2012; Stockemer and Parent, 2014).[1] Our contribution to explaining at least some part of this ambiguity focuses on problems arising in the interpretation of time series relationships; causality among variables that are correlated over time can be misinterpreted and linear insignificance can often hide the presence of a more complicated and meaningful relationship.

In what follows we incorporate income and consumption inequality into an empirical model where voter turnout is viewed as the aggregate outcome of choices made by registered voters whether or not to participate in an election.[2] By using this definition, we emphasize two interrelated parts to the individual's decision to vote: first, the choice whether or not to participate in the election process and, second, the choice over which candidate to support once the decision to vote has been made (instrumental voting).[3] Our model of why individuals choose to participate in election follows Riker and Ordeshook (1968) in emphasizing non-instrumental reasons for voting and is rationalized by analogy with gambling. That is, like instrumental voting the private return from gambling is objectively negative. Yet despite recognizing that there is a negative payoff to gambling, many still choose to gamble. Moreover, while these individuals could simply watch others perform, their enjoyment of gambling comes

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[1] While recent work tends to find a negative relationship arising between turnout and inequality (Ritter and Solt,2019), it remains unclear whether that finding is specific to the time-period chosen and/or the methods used to test the hypotheses. Filette's overview (2016, p.72) is that "empirical indications diverge and give rise to competing theoretical arguments to be tested."

[2] Voter turnout is sometimes defined as the proportion of the eligible population who vote rather than the proportion of those registered who vote, in part because there is often more reliable historical data on the population proportion eligible to vote than on voter registrations (Vowles, 2010). When enrollment is not automatic, unlike the Canadian and Indian cases, an explanation of voter turnout needs to deal with the two-stage choice of whether to register and then whether to vote given registration.

[3] In many statements of the turnout problem only the second part of this decision is given serious consideration. Posed this way, the problem of voter turnout encounters the Downsian (1957) voting paradox; why would any individual choose to vote when the likelihood of influencing the election outcome in their favour is close to zero?

not from watching but from participation, from experiencing the winning and losing associated with gambling. In the same way we argue that in addition to the benefit of having your preferred candidate win the election, individuals have different reasons and hence place different values on electoral participation. Operationalizing this theory involves finding a set of factors that can explain changes in either the instrumental benefit of voting and/or changes in the desire to participate. Individual changes then aggregate into changes in the pool of registered voters who place value on electoral participation.

The paper proceeds in section 2 by outlining a set of factors from the literature that have been found to have influenced voter turnout and sets out their empirical counterparts. These covariates form the controls for the turnout model in which income inequality is expected to play a role. Section 3 presents the initial empirical form of the model to be tested and the expected signs of the coefficients of the covariates in that test. To assess the role played by income inequality, however, the model must pass through two stages. In the first, the model determining voter turnout must both explain the data and generate evidence consistent with the existence of a long run equilibrium arising among the model's covariates. With evidence of an equilibrium model, we can then ask what the data imply for the relationship between income inequality and voter turnout. Section 4 focuses on the time series issues encountered when a linear version of the model is regressed on annual Canadian data and uses an autoregressive distributed lag model to resolve these concerns for Canada. The uniqueness of the nonlinear relationship found for the effect of income inequality on voter turnout requires robustness testing, leading us in section 5 to extend the model and the quadratic role played by the Gini coefficient to a panel of Indian state elections. The results confirm the nonlinearity of the Gini's effect along with the control predictions of the turnout model. The two countries share the result that the positive effect of inequality on turnout peaks at a Gini coefficient of .34. Section 6 summarizes our conclusions.

2. An aggregate model of voter participation

There is a large literature on the factors that have influenced voter turnout (see, for example, the meta-analyses in Cancela and Geys, 2016 and Stockemer, 2017) with more recent concern with the decline in voter turnout that has been observed across many developed countries

(Hooghe and Kern, 2017). Much of the voter turnout literature is cross sectional in nature, where cross country differences in political institutions are used to explain differences in turnout (Blais, 2006; Eichhorn and Linhart, 2020). Other studies have used longitudinal data on individual behaviour, exploiting variables like age, years of education, occupation, and social class (Smets and van Ham, 2013). Here we focus on a set of factors predicted to have affected turnout that can be represented by aggregate time series variables. We then test the resulting model on annual data from Canada, over the 1976 to 2017 time period. The uniqueness of our findings lead us to apply the model on a larger panel data set that covers a set of elections in 14 Indian state elections over the 1957 to 2018 time period. Both countries have well-developed Westminster parliamentary systems with electoral institutions that have remained largely unchanged over these time periods.[4]

A recent concern in the voter turnout literature that has motivated our work has been the linking of a decline in voter turnout throughout the 1990s with the rapid growth in income inequality and its political connection with the Occupy Movement and the earnings of the top 1%. Despite the general acceptance of this relationship, many of the cross-country studies that have focused on this relationship over a longer time period have found little evidence of a negative relationship arising in Western economies, see for example, Fumagalli and Narciso (2011) and Stockemer and Scruggs (2012). On the other hand, others such as Mahler (2008) and Sealey and Anderson (2015) for Canada, have found suggestive connections by pointing to the positive relationship arising between voter turnout and the scale of party redistributive programs. Specifically Canadian studies also suggest that attitudes towards politicians and politics can be inversely affected by growing income inequality (Perrella et al, 2016) while Nadeau et al (2019) show that greater wealth is associated with greater voter participation. In the most recent work done on voter turnout in Canada, Polacko (2020) tests relative power theory (Goodin and Dryzek, 1980) that predicts a negative relationship with income inequality (as lower income voters disengage from the electoral process) against conflict theory (Meltzer and Richard, 1981) that predicts a positive relationship (as growing inequality stimulates

[4] The time periods reflect the earliest availability of annual observations on the Gini coefficients through the latest time periods on which we have observations on competitiveness and winning margins.

political engagement both for and against change). Using a mixture of individual and aggregate data Polacko finds evidence of a significant negative relationship in Canada's ten provincial and federal elections over the 1985-2015 time period.

The variables we use to form our model of voter turnout begin with Downs' (1957) focus on instrumental voting and its predictions for voter turnout. Perhaps most straightforward is the size of the voting pool and its predicted effect on the probability of being a decisive voter (Hansen et al, 1987; Tavares and Raudla, 2018). While the participation of others might be expected to heighten interest in individual participation, Downsian reasoning argues that the larger the number of voters in a constituency, **Constituency_size**, the lower is the likelihood that any one voter will be decisive and hence the smaller will be the benefit of voting in the upcoming election (Gorecki and Gendzwill, 2020). A second factor decreasing the instrumental value of voting would be an increase in the uncertainty of the electoral outcome. Less certainty means that the expectation of being decisive can be held with less conviction (Tavits 2008). We use the historical volatility of party vote shares (**Volatility**) as a proxy for the uncertainty in the expectation that any party will win the upcoming election.[5]

To these variables we add other aggregate variables that have been associated with making voter participation in the election more interesting/valuable. For example, an increase in the number of political parties competing (**Parties**) in the election will attract potential voters by providing a wider range of choice among party ideologies and/or party policy platforms.[6] By fragmenting the vote, more parties also reduce the votes needed to win a plurality, making the likelihood of being a decisive voter larger. Similarly, the competitiveness or closeness of the party/candidate race within the constituency (**Competition**) makes activity surrounding the election of more interest to the voter while also increasing the probability that the vote could be decisive (Eichhorn and Linhart, 2020). The degree of competitiveness arising within

[5] Pedersen's (1979) measure is our measure of vote share volatility.
[6] Zagorski (2020) argues that too much choice (as in too many parties) may generate greater confusion and ultimately lead to lower voter turnout.

constituencies is measured by a multi-party constituency-based measure of electoral competitiveness (Przeworski and Sprague, 1971).[7]

Changes in a country's demographics are also often argued to have mattered in relation to voter turnout. For example, younger age cohorts are seen as less interested and engaged with the political process and so participate with lower frequency in elections than do individuals in older age cohorts (LeDuc and Pammett, 2014; and Dassonneville, 2017). The effect of changes in the relative size of the younger voting pool on voter turnout is tested for in Canada by using the proportion of eligible voters who are between the ages of 18 and 25 (**Youth**). Because the minimum voting age in India was lowered in 1988 from 21 to 18, in the middle of our time period, the effect of demographic change was tested for using the percentage of the population 60 or older (**Old60**). The effect of the discrete change in voting age was incorporated through a dummy variable (**voting_age_18** = 1 in 1989 and later, 0 earlier).[8]

Whether or not a particular candidate is likely to win, the decision to participate may simply reflect voters' discontent with current economic circumstances. That is, disappointing outcomes over a governing tenure may be a potent driver of participation through voters' wishes to reflect their disapproval publicly. If voters are more concerned with economic downturns than upturns, worsening conditions would generate more political dissatisfaction and hence elicit greater participatory response.[9] Here we focus on two aggregate indicators of worsening conditions: increases in the unemployment rate (**Urate**) and/or a fall in the rate of growth of per capita income (**growth_real_inc_pc**) should decrease support for the party in power and increase the demand for alternative programs and parties.[10]

The primary focus of our paper is on the role of income or consumption inequality.[11] While an increase in income inequality can be expected to increase discord as economic differences

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[7] For detail on the derivation of Przeworski and Sprague Index see Ferris et al (2016) and Dash et al (2019).

[8] In Canada the voting age was lowered from 21 to 18 in 1970 (prior to our time period) and has since remained fixed.

[9] A number of writers have found that voters respond more to bad versus good economic outcomes. See, for example, Nannestad and Paldam (1997), Dassonneville and Lewis-Beck (2014) and Dash and Ferris (2020).

[10] The former has been found more effective in Canada (Nadeau and Blais, 1993) whereas the absence of unemployment data for the Indian states led us to use the growth rate of state per capita income.

[11] The absence of information on individual incomes in India led to the use of consumption data for an inequality coefficeient. See the Data Appendix for greater detail.

among voters increase, whether that dissatisfaction translates into voters' political alienation and electoral withdrawal or alternatively stimulates greater political activism and electoral involvement through a desire to promote (and oppose) change through the electoral process has been found to be somewhat problematic (Horn, 2011; Stockemer and Parent, 2014). The different types of response have been addressed in the literature as the difference between relative power and conflict theory. In the empirical work below the Gini Coefficient (based on total income, **Gini**) is used as our measure of economic inequality for Canada and note that Polacko (2020) has presented recent evidence for Canada (1984 and 2015) that higher **Gini's** are associated with a decline in voter turnout and hence supportive of the relative power theory.

Because the pattern of co-movement will feature prominently in the empirical analysis, we conclude this section by presenting in Figure 1 voter turnout in Canada's federal elections between 1976 and 2017, together with Canada's Gini coefficient. In it the widespread recent pattern of turnout decline can be seen, in Canada's case falling from roughly 75 percent of those registered to vote in the 1980's and early 1990's to roughly 60 percent throughout in the 2000's before rising back into the mid 60's in the most recent election (2019). Inspection of Figure 1 also indicates that voter turnout and the Gini coefficient vary in opposite directions for at least part of this time period, providing initial support for Polacko's (2020) finding that increases in the dispersion of income (greater income inequality) have alienated rather than provoked voter participation.

-- Figure 1 about here –

In what follows we will argue that accounting for time series issues reveals that the link between the Gini coefficient and a model of voter turnout is more complicated than the diagram suggests. In particular, the inability to find a linear cointegrating relationship for voter turnout that includes the **Gini** inequality raises the possibility that the observed relationship with turnout is either spurious or nonlinear. Allowing for the possibility that the Gini enters the long run model quadratically raises the possibility that the relationship may embody elements of both relative power and conflict theories.

3. Canadian time series data and initial cointegration tests

The Canadian electoral data used in this paper were collected from the Elections Canada website while the data on the unemployment rate, the Gini coefficient (based on total income) and other economic and demographic variables were taken from the Statistics Canada database, Cansim II. Annual observations for the election variables were constructed by viewing election outcomes as point observations of an ongoing political process and so were interpolated between election years. The remaining variables are available annually. The descriptive statistics of the variables used in our tests are presented in Table 1 along with their time series characteristics.[12] What is important to note is that most of the time series variables are integrated of order one, I(1), meaning that the levels of the variables trend stochastically and become stationary, I(0), only when first differenced. In addition, two of the Canadian variables, the proportion of eligible voters in the 18 to 25 age group and the number of political parties are I(2).

-- Table 1 about here –

The discussion of section 2 suggests the model advanced to explain voter turnout could be tested by estimating an ordinary least squares (OLS) regression that relates voter turnout linearly to the variable proxies. Interpretation of these results is problematic, however, for two reasons: not all of the time series are integrated of the same order and, second, none of the time series for Canada are stationary. To interpret the estimated coefficients of a regression of nonstationary variables as evidence of a long run equilibrium relationship (as cointegrated relationship), all variables need to be I(1) or lower and the residuals of the OLS regression need to be stationary. The first of these problems can be solved by first differencing the two times series that are I(2)--the number of parties and the proportion of young voters (between 18 and 25). Hence the OLS regression test for a long run relationship between voter turnout and the proposed explanatory variables uses the change in the number of parties, **D(Parties)**, and the change in the proportion of eligible voters who are young, **D(Youth)**. Voter turnout is predicted

[12] The Canadian and Indian State datasets are available online at Carleton University’s Dataverse site (Ferris, 2021).

to be larger the larger is the increase in the number of parties and the larger is the fall in the proportion of voters who are young. This regression test estimates by ordinary least squares,

$$Voter\ turnout = \alpha_0 + \alpha_1 Size + \alpha_2 Volatility + \alpha_3 Urate + \alpha_4 Competitiveness + \alpha_5 D(Youth) + \alpha_6 D(Parties) + \alpha_7 Gini + \varepsilon, \quad (1)$$

where α_1, α_2 and α_5 are expected to be negative in sign while α_3, α_4 and α_6 are expected to be positive. Finding $\alpha_7 > 0$ (< 0) would be consistent with conflict (relative power) theory while ε is a white noise random variable. The results of this OLS regression are presented in Table 2 below where the two differenced variables are introduced separately before being used in combination.

-- Table 2 about here --

The results in Table 2 appear broadly consistent with the hypotheses advanced to explain voter turnout suggesting that the model of voter turnout works well. All 16 of the coefficients of the control variables in the three versions of the test in columns (1) through (3) have their predicted sign and with the Gini coefficient, all three regressions explain roughly ninety percent of the variation in voter turnout in our 1976-2015 time period. Moreover, of the 16 predicted coefficient estimates in the three tests, only 2 are insignificantly different from zero. That is, the only prediction that does not receive significant support from the data was the prediction that unemployment rates would be positively related to voter turnout rates. To highlight some of the relationships that appear to be significant, constituency size, party vote volatility at the national level, the relative weight of younger voters in the voting pool are all inversely related to voter turnout as predicted. Greater political competition as within electoral constituencies and increases in the number of political parties are both associated significantly with increases in voter turnout.

Because voter turnout is bounded between 0 and 1 with variation within a narrow range, the model of column (3) was re-estimated using a fraction logit model as a robustness test. The marginal effects of that model are presented in column (4). The similarity in size and significance of the coefficient estimates in columns (3) and (4) suggest that boundary considerations have not presented a serious issue.

Finally, in all forms of the test the coefficient estimate for the Gini coefficient is found to be negative and significantly so. Assuming that the Gini coefficient is exogenous in relation to the turnout model, this result suggests that the longer run rise in the Gini coefficient in Canada has been a significant determinant of Canada's fall in voter turnout. In short, the OLS results appear to reflect a highly successful model of voter turnout in relation to which the relevance of the relative power theory of income inequality is confirmed thus supporting the recent findings and conclusions of Polacko (2020).

As is well known, however, OLS regressions of nonstationary variables that exhibit high R^2s and low Durbin Watson statistics (as is the case here) often produce spurious correlations. While standard errors may suggest the presence of significant causal relationships, these relationships often arise from separate independent responses to a common stochastic trend.[13] The test for whether the set of covariates is consistent with a long run equilibrium relationship amongst the variables (whether the variables are cointegrated) is whether the residuals of the equation are stationary. In Table 2 the augmented Dickey Fuller (ADF) test statistics for all three versions of the voter turnout test, presented at the bottom of Table 2, are not large enough in absolute size to allow rejection of the presence of unit root in the residuals of an equation with six or seven covariates. It follows that some of the coefficient standard errors are likely to be biased so that the seeming strong results of Table 2 and the variables appearing to be highly significant must be viewed with some degree of caution. At least some of the hypotheses that appear to be confirmed by the data are likely to be spurious.

4. Dynamic stability and nonlinearity: An ARDL model of voter turnout

To handle time series issues the autoregressive distributed lag approach (hereafter ARDL) of Pesaran, Shin and Smith (2001) has proven to be particularly useful. It is designed specifically to assess whether or not a cointegration relationship exists among a group of variables when the orders of integration are ambiguous and the sample size is small. If cointegration is found (as indicated by the Bounds test), the ARDL method generates not only the long run equilibrium

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[13] For example, one such spurious correlation produced the hypothesis that the shorter the length of women's skirts the higher the stock market.

time path but also the short run convergent process that surrounds the long run. In generating the dynamic processes that produce this outcome, the method adopts the lag structure that best captures the varying degrees of persistence exercised by each of the interrelated variables.

An ARDL version of our voter turnout model can be written as:

$$Voter\ turnout_t = \alpha + \sum_{i=1}^{i=3} \gamma_i Voter\ turnout_{t-i} + \sum_{i=0}^{i=3} \delta_{ji} Z_{jt-i} + \epsilon_t, \qquad (2)$$

where the Z_j are the j = 7 covariables from equation (1). All covariates are allowed to have up to three lagged terms and ϵ_t is a white noise random variable. The ARDL model then assesses all combinations of current and lagged terms and generates the coefficient estimates that provide the best fit with the data.

Table 3 presents in columns (1) and (2), two error correction versions of the ARDL model described in equation (2). Column (1) includes the full set of covariates used in the OLS regressions of Table 2 while column (2) drops the change in the proportion of voting pool that is young, D(Youth) (whose long run coefficient was found to be insignificantly different from zero in column (1). The reason for retaining the Gini coefficients in the models of columns (1) and (2) will become apparent below.

--insert Table 3 about here –

The results in columns (1) and (2) are quite striking in their own right and in relation to the results found in Table 2. First, as a description of a stable long run relationship connecting voter turnout with constituency competitiveness, party vote volatility, changes in the number of participating parties, constituency voter size and the unemployment rate the model is quite successful. The full error correction model explains close to ninety percent of the variation in voter turnout about the equilibrium time path (with an adjusted R^2 of .891). Unlike the residual ADF tests for cointegration in the OLS models of Table 2, the Bounds test, presented in the bottom line of each column, generates values that clearly reject the hypothesis that there is no relationship among the covariates in levels and thus are consistent with the presence of a cointegrating relationship among the model's I(1) variables. The Canadian data then provide evidence of a long run equilibrium time path among the model's variables towards which short run departures from the path converge through time.

By confirming the presence of a cointegrated long run relationship among the control variables of our model, our confidence that the standard errors generated under the ARDL model are unbiased is increased. As such the new results associated with the individual hypotheses in Table 3 call for some revision of the perspective suggested by Table 2.[14] Perhaps unsurprisingly, while most of the hypotheses that appeared to be confirmed by the OLS form of the test are re-affirmed by the results of the ARDL model, two relationships that appeared to be significant are revealed to be statistically insignificant and one that appeared insignificant is now found to be highly significant.

The hypotheses whose sign prediction remain consistent with the data include voter turnout rising with the average degree of political competition at the constituency level and the number of political parties and falling with a rise in both vote volatility and constituency size. In addition, the predicted positive effect of unemployment on voter turnout that was found to be insignificant in the earlier OLS test is now revealed to be significantly positive. The data is then consistent with the hypothesis that voter discontentment with worsening economic outcomes (as represented by a higher unemployment rate) will result in higher voter turnout.

On the other hand, the coefficient of D(Youth) that was both large in absolute size and significantly negative in the OLS regressions is now revealed to have no significant long run impact on voter turnout in our time period. Similarly, the Gini coefficient that was also indicated as being large in absolute size and highly significant is now revealed to be both small and insignificant. That is, while the earlier OLS results suggest that increases in both these factors played a prominent role in discouraging voter participation, the ARDL results imply that these implied results were spurious.

The fall in the significance of the two covariates that were highly significant in the OLS equations in combination with the continuing significance of the Gini coefficient in the short run led us to investigate whether the imposed linearity of the ARDL model could account for the loss of their significance. In Figure 2 we present the result of running a nonparametric regression examining the possibility of a nonlinear relationship arising between the Gini coefficient and voter turnout.[15] As

[14] We note that the existence of a cointegrating vector does not imply that the sign and significance of the model's coefficient estimates capture the 'true' degree of causality running from the right hand side covariates and voter turnout. Hence while at best we can argue that the sign and significance of covariate coefficients are consistent with the prediction of a particular theory, the test for the existence of a long run equilibrium does not require the absence of endogeneity arising between any subset of the covariates.

[15] Similar modelling of D(Youth) produced no discernable relationship with voter turnout.

that diagram illustrates, there appears to be an inverted U-shaped relationship between the two with the effect of the Gini on voter turnout peaking at about .32 in the lower range of Gini values. That is, the nonparametric relationship suggests that increases in income inequality are at first associated with increases in voter turnout but as the degree of income inequality keeps rising, the effect of greater inequality falls on the margin until a tipping point is reached after which further increases decrease voter turnout. Beyond the peak, additional inequality leads to further voter alienation and turnout decline.[16]

To test whether the Gini coefficient should be present at all in the cointegrated model we first used the lasso linear model selection procedure (Tibshirani, 1996) to confirm the importance of the Gini coefficient to the model's predictive power. Then to assess whether a nonlinear representation would improve the predictive power of the model over its absence (as suggested by columns (1) and (2)) we reran the ARDL model first excluding the Gini coefficient entirely (in column (3) of Table 3) before including the Gini coefficient quadratically (in column (4)). The results in columns (3) and (4) indicate that the quadratic representation does indeed substantially increase the model's fit with the data (the adjusted R^2 rising from .867 to .918) while leaving the other coefficient estimates largely unchanged from those estimated in columns (1) and (2).[17] The results confirm the inverted U-shaped relationship shown in Figure 2 and imply that the effect of the Gini on voter turnout peaks at a value of .34, very close to the peak of .32 indicated in Figure 2. The confidence intervals about the tipping point are presented at the bottom of column (4).[18] Interpreting this result in terms of Figure 1, the data suggest that prior to the 1990's increases in income inequality were

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[16] Doing the same thing for D(Youth) produced no easily categorizable pattern. This led us to accept the finding of its insignificance for our time period and so exclude D(Youth) as a covariate in the final equations of Table 3.

[17] An F test, F= 16.82, allows us to reject the null hypothesis that the added variables do not add to the explanatory power of the model.

[18] Because the tipping point is the ratio of two regression parameters, a number of statistical issues arise when trying to establish an appropriate confidence interval. To do so we adopt two strategies. First, the Delta method uses a truncated Taylor series expansion and asymptotic theory to establish a Wald-type confidence interval. Adopting the 95% confidence criterion establishes Delta lower and upper bounds of .330 and .352. An alternative method that does not require strong identification and hence is robust to mistakes in modeling the form of the relation uses Fieller (1954). Using the ninety five percent confidence criterion, the Fieller method produces lower and upper bounds about the .342 estimate of .334 and .422. While the Delta method provides better tightness about the estimated tipping point the Fieller method better captures the asymmetry witnessed in the nonparametric diagram. See Ferris and Voia (2015, Appendix B.2, pp 183-4) for a detailed description of how these two sets of measures are constructed.

associated with (ever smaller) increases in voter turnout whereas further increases running into the late 1990's and through 2018 were associated with growing voter alienation with the political process and a decline in turnout.

While the long run cointegrating process is the primary point of interest, the short run and error correction processes described by equation (3) are also of interest. The error correction coefficient (-.762 in column (4)) is relatively large in size indicating that random shocks and covariate changes that produce departures from the long run equilibrium time path are returned quickly.[19] And while short run process overall is quickly convergent, the short run effects arising from the various covariate changes have a widely varying set of adjustment processes. For example, an increase in political competitiveness within the constituency has a large short run effect on voter turnout that falls off quickly through time while an increase national vote volatility is associated with an oscillating downward movement in turnout as it adjusts to its lower long run value. Finally, the marginal effect produced by a further increase in income inequality increases in the size of its effect over time, reinforcing the inverted U-shape of its long run relationship with voter turnout.

5. The case of Indian States

Our finding in the Canadian data that there is an inverted U-shaped relationship between the Gini coefficient and voter turnout is significant because it suggests that the empirical ambiguity often found in the relationship between income inequality and voter turnout may not mean that the effect on turnout can be attributed to either the relative power or conflict theory but that each theory may be applicable to a different range of the income inequality spectrum. However, the relatively short time period over which we have Gini observations in Canada raises the question of whether the inverted U-shaped relationship found is unique to Canada and/or our sample period and hence whether this result will generalize across other countries. Because of earlier work done with India, a country with a similar Westminster parliamentary heritage to Canada, we have access to data on a multiplicity of elections across Indian states taking place between 1957 and 2018. We

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[19] The after-tax Gini coefficient moves in parallel with the one based on total income but is roughly 10% lower in size. Hence using the after-tax Gini instead of the total income Gini results replicates the results through Table 2 and the first two columns of Table 3. When the after-tax Gini coefficient is included quadratically the same pattern of coefficient estimates arises, however, the error correction process becomes unstable, oscillating explosively about the estimated long run time path. We speculate that the greater stability of the model with total income is attributable to the greater visibility of pre-tax income to voters.

then use this data as a robustness test of the participation hypothesis outlined earlier on that panel of state data.

The Indian panel data used in our test cover 14 major Indian states: Andhra Pradesh, Bihar, Gujarat, Haryana, Karnataka, Kerala, Madhya Pradesh, Maharashtra, Orissa, Punjab, Rajasthan, Tamil Nadu, Uttar Pradesh, and West Bengal.[20] The panel consists of a maximum of 195 state elections held between 1957 and 2018. The economic and demographic variables applicable to each election were obtained from the *Central Statistical Organization* and *Census of India* while Indian election data was collected from the Electoral Commission of India's website (http://eci.nic.in/eci/eci.html). The database itself is available online at Ferris (2021). In Table 4 we present the descriptive statistics and, when these are compared to their Canadian equivalents in Table 1, it is apparent that the Indian states exhibit considerably more variation in the scale and distribution of constituency sizes, per capita income growth, number of political parties and Gini coefficients.

-- Table 4 about here --

Applying our voter turnout model to the Indian data requires some modification of the form of the test both for data availability reasons and to incorporate electoral features specific to India. In terms of the former, state unemployment rates are unavailable in India so that the hypothesis that voters register disapproval with the level of economic performance arising under the incumbent government is now tested for using the average growth rate of state per capita income over the incumbent's governing period. Voter turnout is expected to be inversely related to the growth rate. Similarly, individual income data is unavailable in India so that our Gini coefficient was constructed from information on consumption expenditures at the household level extracted from nationally representative periodical surveys.[21] The degree of political competition in constituency elections is now proxied by the average size of the winning

[20] These 14 states cover about 85 percent of the Indian population and exclude states such as Assam, and Jammu and Kashmir that were subject to insurgency and other forms of electoral violence. See Diwakar (2008) for a complementary analysis covering all Indian States using macro level data and Panda (2019) who uses political-economic data to analyze voter turnout using individual level data.
[21] See the Data Appendix for detail on how the Gini coefficients were constructed from consumption data.

vote share margin across state election constituencies. In this instance, a larger winning margin indicates less competition and hence is predicted to generate lower voter turnout. The prediction that voter turnout will be lower (larger) the larger the proportion of the population that is young (old) is tested for in two ways. First, the 61st Amendment to the Indian Constitution lowered the voting age from 21 to 18 for all elections in 1989 and beyond. The expected negative effect this would have on the proportion of registered voters was tested for with a dummy variable (1 in 1989 and thereafter, 0 before). We also used the percentage of the population older than sixty to test the prediction that older voters are more engaged in the political process and so are more likely to vote.[22] The Indian Constitution (Article 356) also gives the central government the right to impose president's rule in cases when there is a perceived failure of democratic government (often due to the inability of state parties to form a majority government). In such cases the state is governed directly by the central government in the form of an appointed governor. Periods of president's rule are then periods where state voters are disenfranchised and can be expected to be less interested in electoral participation and turnout.[23]

--Table 5 about here --

Our test of the voter participation model on Indian states appears in Table 5 as a series of dynamic fixed effects models, where voter turnout in the state's previous election was used to account for persistence in turnout over time.[24] In column (1) we present the linear version of the fixed effects version of the model that corresponds to equation (1) introduced earlier. The results indicate that the linear fixed effects model overall explains somewhat over sixty percent of the variation in voter turnout as well as exhibiting some degree of voter turnout persistence

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[22] Data on state-wise percentage of older population is provided in the Census of India. To extend the data from 2011 to 2018, we used data provided by the World Bank. The World Bank, however, provides data only for the national level. By assuming that the difference between national and state figures in the 2011 census persists, we generated figures for each state in 2018. Figures for non-census years were generated by interpolating between census years.

[23] There is one important data outlier corresponding to the 1992 election in the Punjab. This election was the first following the imposition of president's rule imposed to counter a Sikh extremist insurgency. The voter turnout in this election was exceptionally low (24% versus an average of about 65%). A dummy variable for this election was used to minimize the effect of this election on the other results.

[24] The number of useable observations (179) is insufficient for us to apply ARDL modeling to our panel of Indian States.

across elections. On the other hand, the sign and significance of the individual coefficient estimates support only some of the individual hypothesis predictions. As expected, voter turnout falls with increases in the average growth rate of real income per capita over the incumbent's governing term, higher vote volatility among parties, the presence of president's rule and increases in the competitiveness of electoral races at the constituency level. On the other hand, the linear model does not support the predictions that voter turnout should fall with increases in the voting size of the electoral constituency, the number of political parties nor a rise in the proportion of voters over 60 and the lowering of the voting age from 21 to 18 in 1988. As was true of the Gini coefficient in the linear version of the Canadian model discussed earlier, there is no evidence of a role being played by inequality in the linear version of the Indian state model. In this case the coefficient estimate is positive but insignificantly different from zero.

The possibility that an insignificant linear coefficient may hide a nonlinear relationship arising between the Gini coefficient and voter turnout, as was the case for Canada, is again suggested by the result of regressing the binary relationship nonparametrically, the results of which are shown in Figure 3. While less dramatic than the corresponding figure for Canada in Figure 2, the diagram gives support to the hypothesis that an inverted U-shaped relationship is present.

--inset Figure 3 about here --

We then test the hypothesis that low levels of consumption inequality stimulate voter's political engagement and electoral turnout while successively higher levels of inequality lead to growing political alienation and corresponding decrease in voter turnout by entering the Gini coefficient quadratically into the voter turnout model. In addition to including the Gini coefficient quadratically, we include constituency size quadratically. This allows us to assess whether evidence of a Downsian voter turnout effect (rejected in the linear version of the model) reappears on the margin, that is, whether the falling likelihood of individual influence on the electoral outcome effects voter turnout on the extended margin rather than overall. The results are presented in Table 5 as columns (2) and (3). Column (3) adds election period fixed effects to the state fixed effects models of columns (1) and (2).

As the results in both columns indicate, the inclusion of these two quadratic terms provides evidence that is consistent with the predicted effect of voter turnout on marginal increases in constituency size and consumption inequality. In both cases the data imply an inverted U-shaped effect. In the case of the voting size of constituencies, the data implies that increases in constituency size first increase voter participation with ever further increases discouraging voter turnout at the margin. Similarly, increases in consumption inequality are associated initially with increases in voter turnout before the negative effect of further marginal increases overcome the initial advantage to decrease turnout. The Gini coefficient estimates in columns (2) and (3) both imply that the tipping point for the Gini's positive effect on voter turnout occurs at .34 a point that is slightly higher than the state average of .32.[25] There is then in the case of Indian states supporting evidence for the hypothesis that conflict theory may explain changes in voter turnout at relatively low levels of consumption inequality whereas the relative power hypothesis becomes increasingly relevant as consumption inequality rises. It is interesting to note that in both the Canadian and Indian state cases, the effect of inequality on turnout peaks at .34.

Whether or not the quadratic forms used for constituency size and the Gini coefficient, the results across the table for the control variables show remarkable consistency in sign and significance. That is, throughout the data indicate that voter turnout is negatively associated with the average degree of vote share volatility within constituencies, the size of the winning party's vote margin, the higher the rate of per capita income growth over the incumbent's governing tenure and with the presence of presidential rule. An increase in the number of political parties is associated positively with higher voter participation and turnout, but only weakly so. The only hypothesis not supported by the data at the aggregate level is the hypothesis that the age composition of the voting pool matters in predicting turnout. In our case neither the proportion of the population above 60 nor the extension of the voting

[25] Using our two methods to establish the 95% confidence intervals about the tipping point, we find that the Delta method generates lower and upper bounds of, respectively, .323 and .362 while the corresponding Fieller method results are .241 and .494. Once again the Delta method implies relatively more precision in the estimate of the tipping point while the Fieller method captures the asymmetry about the peak and better reflects the greater flatness in the nonparametric estimate of the relationship shown in Figure 3.

franchise to individuals between 18 and 21 were found to have any significant influence on voter turnout. In finding the insignificance of voting age demographics the Indian state results echo our earlier findings for Canada.

6. Conclusion

In this paper, we have used aggregate data to test a model of voter turnout first on annual time series data for Canada over the 1976 to 2017 time period, and then on a panel of 14 Indian state elections held between 1957 and 2018. The results are consistent with at least five of the seven individual hypotheses: voter turnout increasing with an increase with the number of competing parties, the competitiveness of constituency elections and the unemployment rate; while decreasing with greater party instability (higher vote volatility) and the average voting size of an election constituency (marginal size in the case of Indian states). In the Canadian case, the set of hypotheses explain ninety percent of the variation in voter turnout while in the case of the Indian states roughly sixty percent of the variation is explained.

Of the remaining two hypotheses, perhaps the most surprising finding has been that population demographics have produced no significant effect on voter turnout. In Canada the proportion of voters between 18 and 25 had no significant effect on voter turnout and in Indian state elections, neither the proportion of voters above 60 nor the discrete extension of the franchise to voters between 18 and 21 in 1989 had a significant effect. Given its prominence in the literature, some effect on voter participation and turnout might have been expected from the continuous fall in the size of the youth vote that has taken place over our time period (from 19.9% to 11.2%) in Canada and/or the rise in the share of the older electorate in India (from roughly 5% to 10%). However, that has not been the case.

The key focus of our analysis has been the relationship between the Gini coefficient and voter turnout. Increasing voter alienation has often been used to explain the recent decline in voter turnout and the rise in inequality is sometimes suggested as an explanation for why voters have become more alienated. We find that while a multivariate OLS regression appears to support the voter alienation hypothesis, a correction for the time series issues encountered suggests that this seemingly significant result is spurious. Further investigation of the voter turnout/Gini coefficient relationship finds the relationship to be non-linear with voter turnout in Canada taking an inverted

U-shaped form with income inequality peaking in its positive effect on voter turnout at a Gini value of .34. The relatively small number of observations in the Canadian case led to the extension of the analysis to a panel of Indian state elections. Once again when the Gini coefficient was entered linearly, the effect of income inequality on voter turnout was found to be insignificant. Its entry quadratically was found to be highly significant while also increasing the significance of the other covariates in the model.

In both the Canadian and in Indian cases the inverted U-shaped form found in the data is consistent with the hypothesis that from low levels of inequality increases in (income or consumption) inequality first generate greater political involvement before leading to voter discouragement with political activism withdrawal from electoral participation. Canada appears to have entered into this later stage around 2000, with the recent decline in the Gini suggesting some reversal of this process of disengagement. In the Indian case, different states currently are positioned on either side of the .34 tipping point, with the states as a whole experiencing a slow upward trend in the Gini over the time period as a whole.

Appendix on Special Data Sources, Figures and Tables:

Canada. Statistics Canada Table 11-10-0134-01 Gini coefficients of total income. Cansim II v96439638 online DOI: https://doi.org/10.25318/1110013401-eng.

Unemployment rate Cansim II v2062815, monthly average.

Population estimates by age groupings Cansim Table 17100005: v466677, v4666680 v4666683, v4666686.

Elections Canada/Elections Current & Past/Voter turnout at Federal Elections and Referendums https://elections.ca/content.aspx?section=ele&dir=turn&document=index&lang=e

India State level Gini coefficients were estimated from household consumption expenditure data collected by The National Sample Survey Office (NSSO) of the Ministry of Statistics and Programme Implementation through periodic nationally representative surveys. Two types of surveys are conducted: quinquennial (or "thick") rounds done at five-year intervals on a large sample of households and annual/semi-annual (or "thin") rounds undertaken during intervening periods on smaller samples. The government has made this household level data available for public use, retrospectively from the thirty-eighth round in 1983. Expenditure information from both thick and thin annual rounds were used to estimate Gini coefficients for the Indian states. To extend the dataset before 1983, we used the Gini coefficients estimated and provided by Ozler, Datt, and Ravallion (1996). Two adjustments were made to make their Gini coefficients compatible with ours: (1) Ozler, Datt, and Ravallion estimated Gini coefficients for rural and urban areas separately; whereas we used rural and urban population weights to generate aggregate Gini coefficients; and (2) official reports were used to provide the number of households in different consumption intervals. There are some disparities between the Gini coefficients estimated from household consumption expenditures and consumption expenditure intervals for the common survey rounds. For most of the states, the pattern of disparity goes in the same direction. We used the disparity ratio to revise the Gini coefficients backward and for years when consumption expenditure information is unavailable, Gini coefficients were generated by interpolating between survey rounds.

Election data: Election commission of India and calculation of authors

Economic data: Central Statistical Organisation (CSO) and Census of India

The datasets used in the tests are available online at Carleton's Dataverse site (see Ferris, 2021).

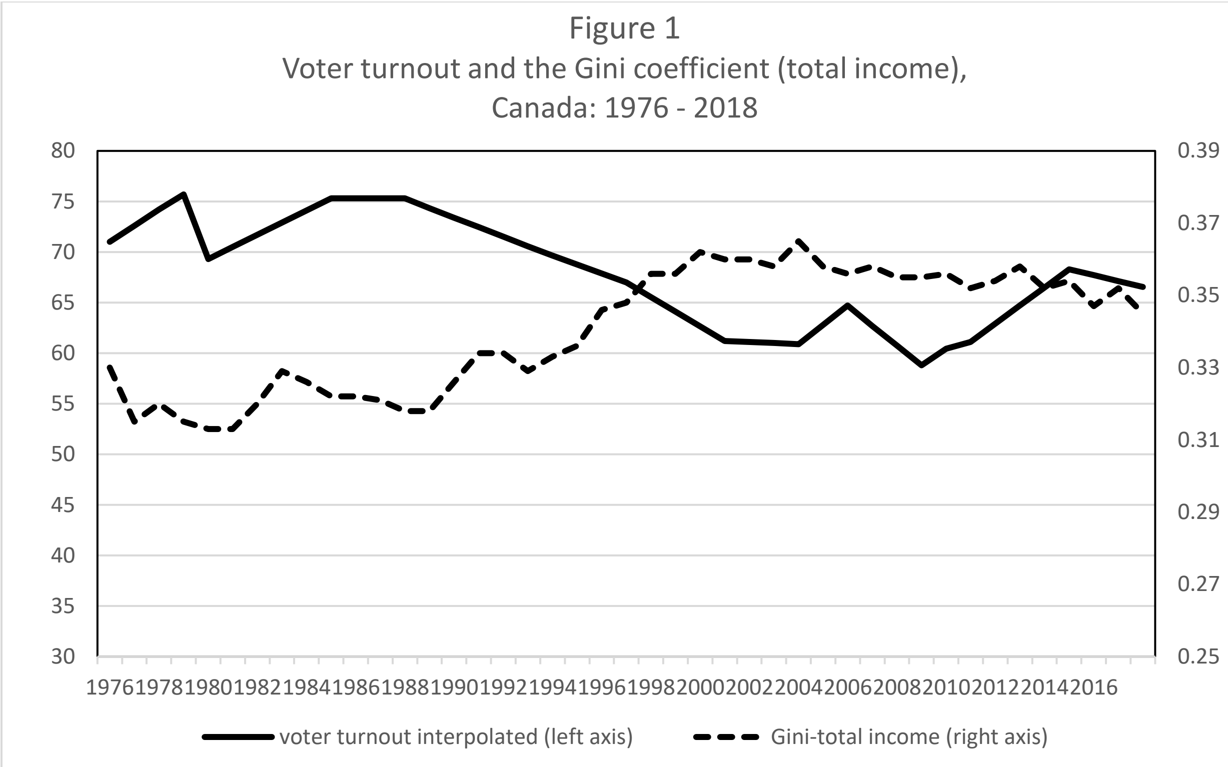

Figure 2: Voter Turnout in Canadian Federal Elections and the Gini Coefficient

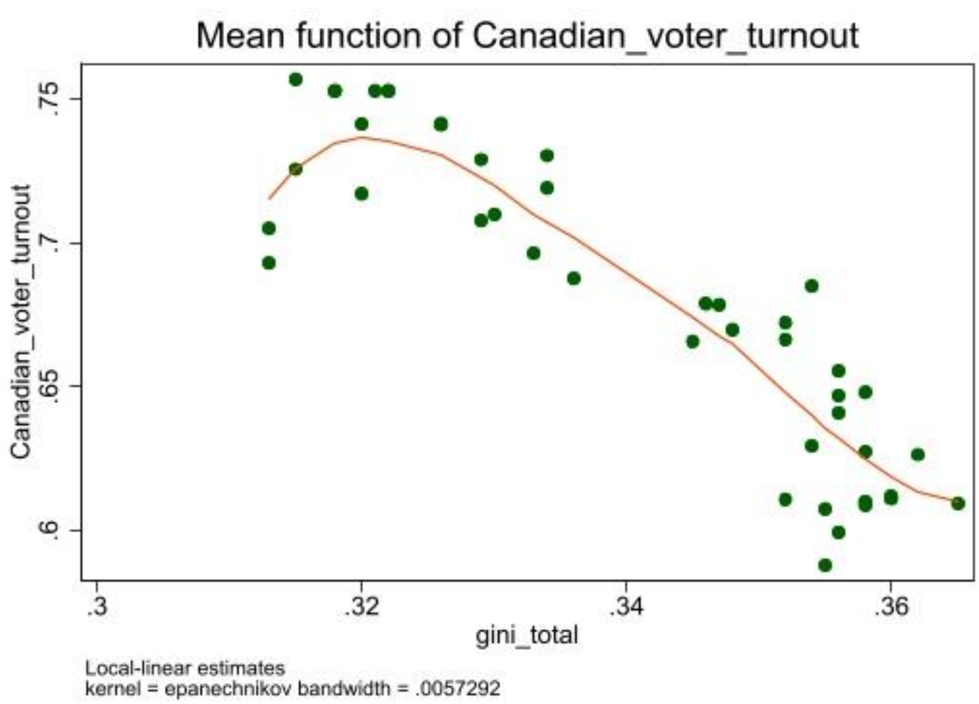

Figure 3: Voter turnout in Indian State Elections and Consumption Gini

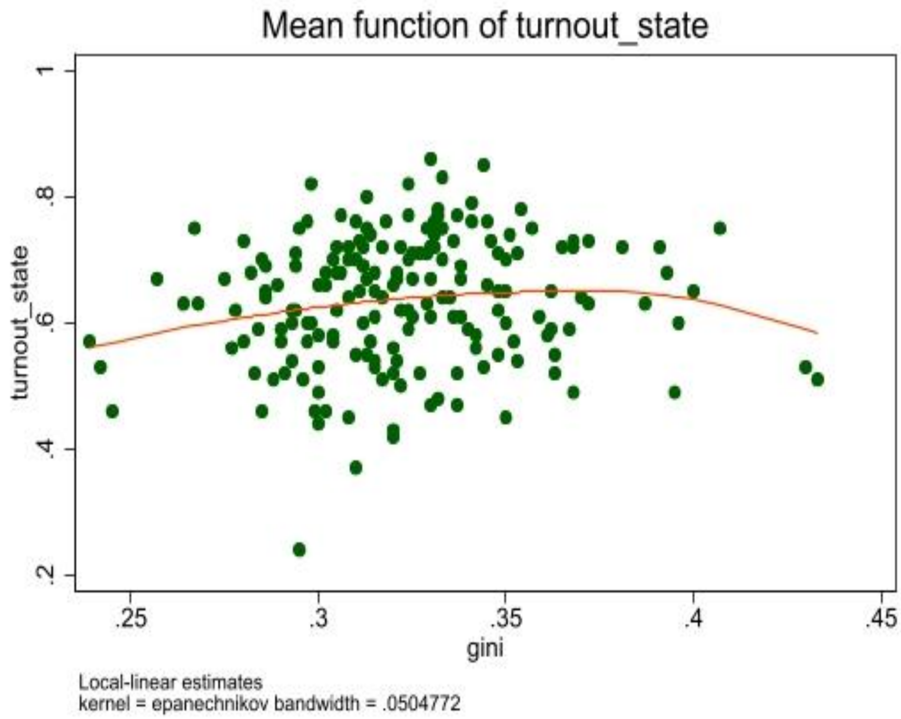

Table 1
Descriptive Statistics: Canada 1976 - 2018

Variable Name	Observations	Mean	Standard Deviation	Minimum	Maximum	ADF - Level -Difference
Voter turnout	44	.679	.053	.588	.757	-1.067 -5.414***
Party Vote volatility at the National level	44	.158	.087	.035	.416	-2.869 -7.935***
Voter Size of Constituency (in 1000s)	44	67.86	8.72	52.3	80.99	-1.138 -3.546**
Unemployment rate	44	8.177	1.65	5.7	12	-1.626 -4.748***
Gini Coefficient (total income)	44	.340	.017	.313	.365	-1.067 -7.731***
Average candidate competitiveness at the constituency level	42	.615	.132	.417	.887	-1.591 -3.705***
Change in the proportion of eligible voters between 18 and 25 years old	44	-0.002	.002	-0.008	.003	-0.993 -3.764***
Change in the number of political parties	44	.332	.630	-1	2	-2.904 -7.935***

*** 1% (** 5%) critical value -3.634 (-2.950)

Table 2

OLS Voter Turnout Regressions: Canada 1976 - 2017

(absolute value of t-statistic using robust standard errors)

Explanatory Variables Predicted sign in brackets		Voter turnout (1)	Voter turnout (2)	Voter turnout (3)	Fractional Logistic Regression Marginal Effects (4)
Voter size of Constituency	(-)	-0.0020**	-0.0023**	-0.0026***	-.0026***
(in 1000s)		(2.50)	(2.27)	(2.75)	(3.16)
Volatility	(-)	-0.171**	-0.255***	-0.182**	-.180***
		(2.58)	(4.27)	(2.85)	(3.18)
Urate	(+)	0.002	0.005*	0.002	.0012
		(0.83)	(1.81)	(0.81)	(0.78)
Competitiveness	(+)	0.176***	0.252***	0.197***	.187***
		(4.54)	(6.92)	(5.15)	(5.54)
D(Youth)	(-)	-5.482***		-5.191***	-6.011***
		(3.42)		(3.39)	(4.19)
D(Parties)	(+)		0.010*	0.007*	.007***
			(2.15)	(2.02)	(2.11)
Gini (Total income)	(?)	-1.333***	-1.529***	-1.076**	-1.022***
		(3.56)	(3.39)	(2.62)	(2.78)
Constant		1.156***	1.197***	1.101***	
		(12.08)	(12.46)	(12.28)	
Regression Statistics					
Observations		42	42	42	42
Adj R^2		.910	.891	.913	
F		76.07	57.67	67.73	Wald (7) 646.26
Durbin Watson		1.09	0.932	1.02	
ADF of residuals: MacKinnon critical value (7) at 10 % = -4.70		-3.77	-3.46	-3.60	

* (**)[***] significantly different from zero at 10 (5) [1] percent.

D(.) indicates the first difference of the variable bracketed.

Table 3

Autoregressive Distributed Lag (ARDL) Models of Voter Turnout (up to 3 lags)

(absolute value of t statistics in brackets)

Explanatory Variables / Predicted sign		D(Voter Turnout) (1)	D(Vote Turnout) (2)	D(Voter Turnout) (3)	D(Voter Turnout) (4)
Lagged voter turnout	(-)	-.687***	-.667***	-621***	-.762***
		(7.69)	(7.97)	(7.34)	(9.90)
Long run cointegration:					
Constituency Size (1000s)	(-)	-.003**	-.003***	-004***	-.004***
		(2.70)	(2.85)	(7.15)	(4.41)
Volatility	(-)	-.856***	-.839***	-.839***	-.776***
		(7.07)	(9.31)	(8.13)	(10.67)
Urate	(+)	.021***	.021***	.015***	.016***
		(5.02)	(6.03)	(5.13)	(4.98)
Electoral Competitiveness	(+)	.582***	.571***	.582***	.526***
		(8.96)	(12.65)	(12.73)	(13.86)
D(Parties)	(+)	.024***	.023***	.031***	.022***
		(3.83)	(4.53)	(8.13)	(5.38)
D(Youth, 18-25)	(-)	.309			
		(0.23)			
Gini (total)	(-)	-.070	-.102		27.50**
		(0.15)	(0.24)		(2.55)
Gini_squared	(-)				-40.23**
					(2.57)
Short Run:					
Electoral Competitiveness:					
D1		-.240***	-.238***		-0.250***
		(4.82)	(4.98)		(5.79)
LD		-.303***	-.303***		.293***
		(7.85)	(8.05)		(8.51)
Volatility:					
D1		.465***	.463***	.340***	-.504***
		(6.39)	(6.60)	(5.51)	(7.79)
LD		.294***	.291***	.220***	.303***
		(5.12)	(5.33)	(4.22)	(6.12)
L2D		.313***	.311***	.298***	-.312***
		(8.00)	(8.30)	(8.48)	(9.19)
D(Parties)					
D1		.001	.001	.0004	.001
		(0.26)	(0.25)	(0.13)	(0.22)
LD		.010***	.010***	.006	.009***
		(3.54)	(3.64)	(2.53)	(3.41)
Gini (total)					
D1		-.223	-.213		
		(0.69)	(0.68)		
LD		-.316	-.280		
		(1.06)	(0.67)		
L2D		-.993***	-.995**		
		(2.94)	(3.03)		
Gini Squared					
D1					-.605

				(1.42)
LD				-1.23*
				(1.93)
L2D				-1.49***
				(3.37)
Urate				
D1	-.011***	-.011***	.007***	-.011***
	(4.85)	(4.99)	(4.22)	(5.45)
LD	-.004**	-.004*	-.003	-.004**
	(2.26)	(2.33)	(1.70)	(2.68)
Constant	.334**	.348***	.357***	-3.09**
	(2.59)	(3.08)	(4.49)	(2.19)
Observations	39	39	39	39
Adj R^2	.891	.891	.867	.918
Bounds Test: No levels relationship				
F (k=7 1% lower bound 3.74 upper 5.06)	20.63***	23.15***	21.81***	28.17***
t (k=7 1% lower bound -3.43 upper -4.60)	-7.69***	7.91***	-7.339	-9.90***
Gini Tipping point				.342
Gini 95% Confidence Intervals:				
Delta (symmetric about peak)				.330 .352
Fieller (asymmetric about peak)				.334 .422

Notes: D1(2) refers to first (second) difference; L1(2), one (two) lags; LD, the lag of the first difference, etc.
* (**)[***] significantly different from zero at 10 (5) [1] percent.

Table 4

Descriptive Statistics for 14 Indian States, 1957-2013

Variable Name	Observations	Mean	Standard Deviation	Minimum	Maximum
State Voter turnout	195	.627	.109	.24	.86
Average constituency vote volatility	180	.380	.130	.12	.72
Constituency Size (in 1000s)	183	134.664	61.65	39.25	371.052
Average growth rate of per capita income over governing term of tenure	181	3.13	3.27	-6.42	16.43
Gini Coefficient	183	.323	.032	.24	.43
Winning Margin	195	.145	.046	.06	.29
Percentage over 60 (Old)	183	7.04	1.40	4.97	13.45
Voting age change 21 to 18 (1988)	195			0 1957-1988	1 1989-2013
Punjab 1992 election dummy	195			0 (otherwise	1 (in 1992)
Number of political parties	194	28.86	34.47	3	302

Table 5

Fixed Effects Models of Voter turnout in 14 Indian States 1957-2018

(absolute value of t-statistic in brackets using robust clustered standard errors)

	State Voter Turnout Linear (1)	State Voter Turnout Quadratic (2)	State Voter Turnout Quadratic-two way (3)
Last election's voter turnout	.358*** (4.02)	.325*** (3.86)	.231** (2.15)
Growth rate of real income per capita over governing tenure	-.004*** (3.58)	-.004*** (4.40)	-.005*** (5.70)
Constituency size (in 1000s)	.0001 (0.42)	.0009** (2.08)	.003*** (3.57)
Constituency size squared		-1.94e-6* (1.87)	-5.44e-6*** (3.24)
Average constituency Vote Volatility	-.103** (2.66)	-.088** (2.11)	-.098** (2.42)
Gini coefficient	.109 (0.81)	4.75*** (4.68)	5.44*** (4.13)
Gini squared		-6.97*** (4.53)	-7.95*** (4.00)
Winning Margin	-.220** (2.48)	-0.227** (2.29)	-.315 (1.53)
Political Parties	.0001 (0.60)	.0003* (2.03)	.0001 (0.39)
Presidential rule	-.047*** (5.02)	-.044*** (5.11)	-.035*** (2.84)
Old (60 plus)	.004 (0.60)	.002 (0.39)	-.003 (0.31)
Lowering of voting age 21 to 18	.013 (1.21)	-.005 (0.44)	-.036 (1.13)
Punjab_1992_election	-.448*** (54.09)	-.438*** (56.69)	
Constant	.433*** (5.67)	-.371* (1.77)	-.506** (2.32)
Observations	179	179	177
R^2 within	.664	.686	.581
R^2 between	.784	.668	.159
R^2 overall	.659	.529	.383
Gini Tipping Point		.341	.342
Gini 95% Confidence interval			
Delta (symmetric)			.323 .362
Fieller (asymmetric)			.241 .494
Fixed effects 1) State Effects	yes	yes	yes
2) Election Effects			yes

* (**)[***] significantly different from zero at 10 (5) and [1] percent.

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