Short- and long-term growth effects of special interest groups in the U.S. states: A dynamic panel error-correction approach

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Abstract

The perception of special interest groups as a serious threat to economic growth has strengthened over the years; however, the vast empirical literature surrounding this claim has produced mixed and inconclusive results. This study re-examines the issue incorporating a potentially important aspect that has generally been ignored by previous studies, namely, the implicit suggestion by some of the theoretical works that the extent and the intensity of the growth effects of special interest groups may differ significantly over different time frames. Specifically, this study uses dynamic panel error-correction methods (Pesaran, Shin, and Smith (1999)) to properly determine whether these effects, if they exist, occur mostly in the short run or the long run based on data from a panel of 48 U.S. states for the years 1975 – 2004. The joint Hausman-type test selected the preferred model, which controls for human capital achievement, initial income, income inequality, and the tax burden. This model produced results which are in sharp contrast to the simple linearly negative or positive findings reported in much of the literature by indicating that special interest groups have significant non-linearly inverted U-shaped long-run effects on growth, and that it takes time (about 8 years) for the full effects to become evident. The results provide evidence that U.S. states face a threshold point below which special interest groups' lobbying and rent-seeking activities boost long-run growth performance but above which they have damaging effects on long-run growth effort. This is confirmed by the Lind and Mehlum (2010) u-test which also suggests that the threshold point is reached when the activities and strength of special interest groups (measured by the percentage of each state's public and private nonagricultural wage and salary employees who are union members, and which varies from 3.8% to 38.7%)) is at the 15.8% level.

I. Introduction

Olson's (1982) theory of institutional sclerosis has undergone considerable scrutiny and is still much debated as researchers try to substantiate its validity and discern its contributions to the political economy of growth (cf., Choi (1983), Weede (1986), and Coates, Heckelman (2003), and Heckelman (2007))[1]. The theory, in essence, maintains that institutional and political stability provide a fertile ground for the development of special interest groups whose lobbying for preferential policies such as subsidies, tariffs, tax loopholes, and competition-limiting regulations divert resources from productive uses and, thus, gradually

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[1] Heckelman (2007) noted that the many studies testing the theory are roughly equally divided between those based on quantitative regression analysis, in which this paper is located methodologically, and those that use the narrative case study approach.

impede technological progress and economic growth[2]. In accord with the theory, many empirical

studies using mainly cross-sectional data from mostly OECD nations and the U.S. States, but

increasingly from developing nations as well, find a significant negative growth effect for special

interest groups (e.g., Olson (1982), Choi (1983), Weede (1984), Lane and Ersson (1986), Vedder

and Galloway (1986), and Crain and Lee (1999)); however, a number of studies have offered

evidence for some dissenting views in that some find only weak or mixed support (e.g., Gray and

Lowery (1988), Lange and Garrett (1985), and Landau (1985); others find no supporting

evidence (e.g., Nardinelli, Wallace, and Warner (1988), Wallis and Oates (1988), Knack (2003),

and Kang and Meernik (2005)); and still others find evidence suggesting that SIGs growth

effects are actually positive (cf., Bejar (2003) and Heckelman and Coates (2003)), while a few

have suggested an inverted U-shape relationship (Choi (1983), Olson, Heckelman, and Coates

(2003), and Heckelman and Coates (2003))[3]. Thus, although the empirical studies have shed

some light on how special interest groups' activity affects economic growth, the issue is far from

being resolved (see Heckelman (2007) for a thorough review). Accordingly, a re-investigation of

the basic issues is warranted. We do so herein by using newer techniques that allow us to

improve the power of statistical tests by exploiting the richness of panel data which have rarely,

if at all, been applied to the analysis of all of the following issues, which have obvious and

important theoretical and policy implications: (i) do special interest groups affect economic

growth and, if so, what is the sign of the revealed relationship? (ii) If these groups do affect

growth, do the effects occur mostly in the short run or the long run? (iii) If long-run effects exist,

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[2] Admittedly, this is a simple rendition of a complicated theory whose many assertions are difficulty to test because they have not been expressed in explicitly testable forms. Thus, as Quiggen (1992) has noted, empirical tests of Olson's theory, including this study, are confined to testing its basic version.
[3] The focus in this paper is on Olsonian special interest groups (e.g., labor unions) which have a redistributive propensity rather than Putnamian groups (e.g., social service groups such as the Rotary and networking groups) (Putnam (2000)) which may have a positive effect on economic growth and well-being.

how quickly are they realized? In other words, how long does it take for the long-run growth effects to become evident following a special interest groups' activity shock? (iv) What is the nature of the revealed relationship between these groups and growth, that is, is the relationship linear or nonlinear? (v) If the relationship is nonlinear, is it an inverted U-shape or a U-shaped relationship and at what level of special interest groups' activity does the turning point occur? This paper provides answers to these questions in the context of the U.S. States for the 1975 to 2004 years.

The next section briefly reviews various views on how special interest groups affect economic growth. Section 3 outlines the econometric methodology and describes the data employed. Section 4 presents the empirical results and Section 5 overviews the contribution of the study.

II. Special Interest Groups and Economic Growth: A Brief Review

Previous empirical studies on the effects of special interest groups (SIGs, henceforth) activity on economic growth have tended to use cross-sectional and other methods which may capture only the long-run growth effects. In other words, these studies have not explicitly separated the short-run from long-run growth effects of special interest groups activity. Yet there are plausible theoretical arguments suggesting that these effects may differ significantly over different time frames. Given that Olson's theory is inherently long-termed (cf., Bejat (2003) and Zimmermann and Horgos (2008)), there may indeed be considerable time lags before shocks in SIGs activity go through their short-run dynamic path to long-run equilibrium when their full effects become evident. As a result, there is much opportunity for different short-run SIGs growth effects scenarios to arise, and in a number of ways. First, as Olson (1982) argued, the

number of special interest groups and their strength and effectiveness are positively related to the amount of time they exist in the stable environment. As a result, SIGs activity effects should rise monotonically with the passage of time and becomes increasingly harmful to economic growth (*"monotonic view"*). Second, Choi (2012) offers another possible source for the temporal difference in SIGs growth effects by suggesting that special interest groups decide on who to lobby (ideologically neutral or strong politicians) and how much money to spend based on whether they care more about short-run or long-run distributive policy outcomes. His findings suggest that these groups are shortsighted in that they are primarily motivated by short-run incentives, implying that their effects on growth may be particularly harmful in the short run (*"myopic view"*). This view is consistent with the prediction of certain political economy models (e.g., Alesina and Tabellini (1990) and Cukierman, Edwards, and Tabellini (1992)), and it can also be gleaned from the work of Caporale and Leirer (2010) who, although they rejected the hypothesis that political instability (measured by the turnover rate in party state governors) increases economic growth in the U.S. States, did find that such instability leads to more rent-seeking activity by governments with a shorter life expectancy and, hence, lower economic growth. In other words, governments with such expectancy tend to focus their efforts on short-run rather than long-run distributive policies. This too suggests that the harm to economic growth takes place mostly in the short run. Third, a contrary view argues that special interest groups are mostly farsighted in their policy preferences because they attempt to maximize their members' expected welfare over the long run (*"farsighted view"*) (cf., Jacobs (2011)). For instance, labor unions are not just concerned about the current wage increases their members receive but also about the security of their pension benefits. If so, then SIGs may aggressively pursue farsighted policy actions that may enhance the groups' future welfare. This suggests that the possible

adverse effects of SIGs on economic growth may be quite significant in the long run than in the short run.

Fourth, Bejar (2003) finds that special interest groups have a significant positive effect on the economic growth of U.S. States and, furthermore, that these groups help close the gap between the growth rates of the poorer and the richer economies of these states (*"the conditional convergence view"*). This gap closes because "….the creation of interest groups in the states with lower initial levels of gross state product promote growth more than in those states with higher initial levels of gross state product" (Bejar (2003. p. 14)). This suggests that the positive growth effects of special groups are relatively higher in the early stages of economic development than in the later stages.

Finally, it has been suggested that the relationship between SIGs activity and growth may be better characterized as non-linear in nature as it may very well be subject to the law of diminishing marginal returns (*"diminishing marginal return view"*) (cf., Choi (1983) and Olson, Heckelman and Coates (2003)). This may hold true if, for example, in the early stages of their formation special interest groups are represented by the most talented and effective lobbyists who will initially have a small but increasingly negative effect on economic growth. However, with time, lobbying effectiveness diminishes due, perhaps, to a limit in the supply of talented lobbyists and, thus, additional SIGs activity will produce less additional harm on economic growth. Thus, this increasing and then decreasing returns to SIGs activity is reflected as an inverted U-shape relationship between growth and such activity.

Heckelman and Coates (2003) offer another reason for an inverted U-shaped relation between growth and SIGs activity. Specifically, they suggest that the nature of the relation

hinges on a nation's level of economic development (*"economic development view"*). Using data from a relatively large sample of countries, their study finds that SIGs have a negative effect on the share of gross domestic product going toward investment in the OECD countries and, hence, lower economic growth; however, the effect is somewhat positive for the developing countries studied suggesting that SIGs increase the share of GDP that is invested due, perhaps, to more SIGs activity causing more social cohesion and social capital (Coates and Heckelman (2003)). Thus, this view, expressed as an explicitly testable proposition, suggests that in the early stages of economic development, the impact of special interest groups on growth is positive, but as development proceeds, this positive impact diminishes and becomes negative in the later stages.

Surprisingly, both the theoretical and empirical literature have given little or no attention to the polar view of the *economic development view,* namely, the possibility of a U-shape relationship between SIGs activity and growth during the process of development. In this case, one may argue that SIGs activity has both growth-retarding (GR) and growth-enhancing (GE) effects, with the GR effects out weighing the GE effects in the early stages of development but with the GE effects dominating the GR effects in the later stages. To offer some motivating hypotheses for a U-shape relationship, we first invoke the following arguments:

(i) The rent-seeking and lobbying activities of SIGs inhibit economic growth by diverting resources from productive uses (Buchanan (1980), Olson (1982), and Murphy, Shleifer, and Vishny (1991)).

(ii) Certain groups (e.g., the rich) have a higher propensity to save than the rest of society, thus, the increase in the relative size of these groups may result in a higher share of output going to saving and investment and, hence, increasing economic growth (cf., Lewis (1954), Kaldor (1957), and Stiglitz (1969)).

(iii)   Much of the benefits from the rent-seeking and lobbying activities of SIGs are likely to accrue to the richer groups in society (e.g., labor unions), resulting in increased income inequality (Knack (2003), Horgos and Zimmermann (2009), and Rossignoli (2013)) which, in turn, enhances growth, based on argument (ii) above. This suggests that SIGs activity affects income distribution (cf., Coates, Heckelman, ans Wilson (2011)). In other words, such activity and income inequality may be inexorably linked over time. This issue is incorporated in our subsequent analysis by including an interactive term involving SIGs activity and income inequality in our growth model.

Based on the above arguments, the growth-retarding (GR) effects of SIGs activity derive from their rent-seeking and lobbying activities, while the growth-enhancing (GE) effects are affected positively by increases in the proportion of the population that is rich and by increases in its propensity to save. These effects can be synthesized to put forth a U-shape relationship between SIGs activity and growth (*the income distribution view*). To see this, consider the simple illustration in Figure 1 in which SIGs activity is on the horizontal axis and growth is on the vertical axis. SIG_T is the threshold SIGs activity level or rate, and $GROWTH_N$ is the economic growth that would exist if SIGs activity were theoretically at zero (SIG_N). With the advent of institutional and political stability, SIGs activity begins to rise above SIG_N. In this early stage, the GR effects are likely to dominate the GE effects because SIGs activity is subject to significant increasing returns (that is, significant growth-deterring effects), as previously argued; however the proportion of the population that is rich, as well as its propensity to save are relatively low (low growth-enhancing effects), resulting in a net negative effects on growth. Note, however, that over time, the increasing returns to SIGs activity will start to diminish (reducing the GR effects), but the increased SIGs activity, as well as the resulting income

inequality will be increasing the proportion of the population that is rich along with its propensity to save (increasing the GE effects). This indicates that growth will be falling at a smaller rate between SIG_N and SIG_T in Figure 1. Eventually, rising SIGs activity will reach the threshold level or rate, SIG_T, with the corresponding economic growth rate being $GROWTH_T$. Further increases in SIGs activity beyond this point will produce positive and increasing growth rates as growth-enhancing forces begin to overwhelm the growth-retarding factors. Thus, this version of the nonlinear hypothesis is based on the notion that in the early stages of institutional and political development, the growth-retarding effects of SIGs activity dominate the growth-enhancing effects, but that the reverse is true in the later stages of development.
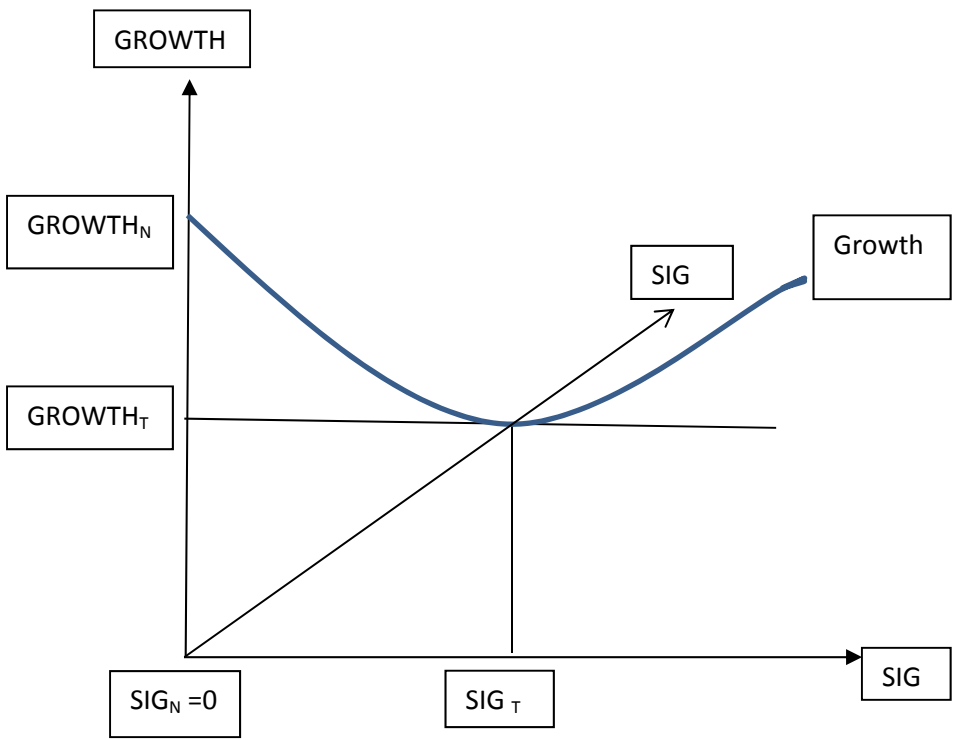

Figure 1: A Temporal relationship between special interest groups and growth

III. Econometric Methodology and the Data

As already stated, the relationship between special interest groups and economic growth is investigated herein by adopting a dynamic panel error-correction approach to properly separate the potential short-run and long-run SIGs activity growth effects. This approach is based on an autoregressive distributed lag (ARDL) model which is estimated using the Pooled Mean Group (PMG) estimator of Pesaran, Shin, and Smith (1999), and a panel consisting of the 48 contiguous U.S. States and spanning the years 1975 to 2004. The use of the PMG offers a number of advantages. For example, it allows consideration of individual state-specific effects that might invalidate the results of, say, cross-sectional methods used in many previous studies, as well as the identification of the long run relationship amongst the variables under study, whether such variables are stationary or non-stationary[4]. In other words, the PMG permits the assumption herein that the long-run relationship between SIGs activity and economic growth rates would be identical across the U.S. states, consistent with the conditional convergence hypothesis (cf., Bejar (2003) and Caporale and Leither (2010)), but that the short-run adjustments to long-run equilibrium differ across the states due to factors such as differences in state moralistic-political subcultures (cf., Thomas and Hrebenar (1999)) and differences in lobbying regulations and the vigor of their enforcement (Hamm, Weber, and Anderson (1994)) and Cole and Chawdhry (2002)).

To investigate the effects of special interest groups on state economic growth, we assume that the long-run relationship between such growth and its determinants can be expressed as:

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[4] As is well-known, many variables are non-stationary in that they have different variances and correlation between past and future values. A stationary variable is said to be integrated of order zero or I(0); a non-stationary variable needing first differencing to be stationary is integrated of order one or I(1); in the case where second differencing is needed for the variable to be stationary then it is integrated of order two or I(2), and so on.

$$y_{it} = \theta_{0i} + \boldsymbol{\theta}^*_{1i}\mathbf{x}_{it} + \upsilon_{it} \tag{1}$$

where the subscripts i = 1,2,…, 48, represent the contiguous U.S. States, and t = 1,2,….T, are for the years 1975 to 2004; y_{it} is the growth rate level of real per capita income; $\mathbf{x}_{it}$ is a vector of control variables including a measure of special interest group activity and its square, all of which are further discussed below;[5] and $\boldsymbol{\theta}^*$ is a vector of parameters.

Assuming that the variables in equation (1) are co-integrated then the error term is an I(0) process for all i (Pesaran and Smith (1999))[6]. Dynamic heterogeneity in the regressions over time and across the states under study can then be introduced by embedding equation (1) in an autoregressive distributed lag (ARDL) of order (p, q):

$$y_{it} = \lambda_{1i}\, y_{it\text{-}1} + \lambda_{2i}y_{it\text{-}2} + \ldots + \lambda_{pi}y_{it\text{-}p} + \delta^*_{0i}x_{it} + \delta^*_{1i}x_{it\text{-}1} + \ldots + \delta^*_{qi}x_{it\text{-}q} + \gamma_i + \epsilon_{it} \tag{2}$$

where γ_i represents the fixed effects, and the other variables are as already defined.

Taking the differences of equation (2) and rearranging terms yield the vector error-correction model:

$$\Delta y_{it} = \sum_{j=1}^{p-1}\gamma_{ji}\Delta y_{it\text{-}j} + \sum_{j=0}^{q-1} \delta^*_{ji}\, \Delta x_{it\text{-}j} + \varphi_i[y_{it\text{-}1} - \theta_{0i} - \theta^*_{1i}x_{it}] + \epsilon_{it} \tag{3}$$

where γ_{ji} and δ^*_{ji} are the short-run coefficients, θ_{0i} and θ^* are the long-run coefficients, and φ_i is the speed of adjustment (error-correction term) of the dependent variables towards its long-run equilibrium given a change in the independent variable. $\varphi_i < 0$ ensures that the long-run relationship exists. Thus a significant and negative value of φ_i indicates co-integration between y_{it} and x_{it}.

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[5] It is customary to use the quadratic method to test for non-linear relationships. However, this method has recently drawn criticism from the important paper by Lind and Mehlum (2010). Thus a more appropriate test suggested by these authors is applied later to the data to test the validity of our results.
[6] The validity of this and related assumptions is tested below using unit root and panel co-integration tests.

For the purposes of our study, the ARDL specification (equation (2)) with a maximum lag of one for each variable is expressed explicitly as[7]:

$$y_{it} = \mu_{it} + \delta_{10i}lnSIG_{it} + \delta_{11i}lnSIG_{it-1} + \delta_{20i}ln(SIG^2_{it}) + \delta_{21i}ln(SIG^2_{it-1}) + \delta_{30i}lnGINI_{it} +$$

$$\delta_{31i}lnGINI_{it-1} + \delta_{40i}ln(GINI_{it} * SIG_{it}) + \delta_{41i}ln(GINI_{it-1} * SIG_{it-1}) + \delta_{50i}lnHUCAP_{it}$$

$$+ \delta_{51i}lnHUCAP_{it-1} + \delta_{60i}lnSLTAX_{it} + \delta_{61i}lnSLTAX_{it-1} + \lambda_i lny_{it-1} + \epsilon_{it} \qquad (4)$$

where, as already noted, y_{it} is the state per capita income, and it was taken from the Regional Accounts Data found at the Bureau of Economic Analysis's website and was deflated using the Bureau of Labor Statistics Consumer Price Index (1982-84 = 100); SIG represents the strength of special interest groups, and it is measured by the percentage of each state's public and private nonagricultural wage and salary employees who are union members, arguably one of the most important special interest groups (cf., Caporale and Leirer (2020))[8]. The square of SIG is included to account for possible nonlinearity, as discussed earlier in the literature review section.

To control for the possible effect of state differences in income inequality on growth the state gini coefficient (GINI) is included.   It is not clear what its impact will be as some studies have suggested a positive relation (Partridge (1997) and (2005)), others have found a negative impact (Frank (2009)) and still others suggest no impact at all (Panizza (2002)). Data for GINI are from Frank (2009) who computed them using tax data reported in *Statistics of Income* published by the IRS.

As already noted in the case of *the income distribution view*, SIGs activity and income inequality (GINI) are likely to be inexorably linked over time, as they influence growth. The interactive term (GINI$_{it}$ *SIG$_{it}$) is included to account for such possible interaction. HUCAP

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[7] The lag of one is imposed given the length of the annual data (30 years). See footnote 9 for related comments.
[8] Data on State union membership are available at http://www.trinity.edu/bhirsch/unionstats/.

proxies human capital investment as measured by the two variables HSCHOOL and COLLEGE, the proportion of the state population with at least high school diploma and the proportion with at least a college degree, respectively, and are taken from the annual Current Population Survey.

Some studies (e.g., Fox and Murray (1990), Modifidi and Stone (1990), and Cole (2000)) have shown that state and local taxes (tax burden) impede sub-national economic growth. However, others (e.g., Due (1961) and Wasylenko (1997)) have noted that firms consider benefits of public services made possible by such taxes when making location decisions. Thus, higher taxes may actually stimulate economic growth. We control for the state's tax structure by including the ratio of state and local tax revenues to state personal income (SLTAX)[9]. Finally, the conditional convergence hypothesis (cf., Barro and Lee (2001)) is tested for and it is reflected by the initial income per capita.

Taking the differences and rearranging the terms of equation (4) yield the error-correction model to be estimated:

$$\Delta \ln y_{it} = \varphi[y_{it-1} - \theta_{0i} - \theta_{1i}\ln SIG_{it} - \theta_{2i}\ln(SIG^2_{it}) - \theta_{3i}\ln GINI_{it} - \theta_{4i}\ln(GINI*SIG_{it}) - \theta_{5i}\ln HUCAP_{it}$$
$$- \theta_{6i}\ln SLTAX_{it}\ ] + \delta_{11i}\Delta \ln SIG_{it} - \delta_{21i}\Delta \ln(SIG^2_{\ it}) - \delta_{31i}\Delta \ln GINI_{it}$$
$$-\ \ \delta_{41i}\Delta \ln(GINI*SIG_{it})\ - \delta_{51i}\Delta \ln HUCAP_{it} - \delta_{61i}\Delta \ln SLTAX_{it}\ + \epsilon_{it} \tag{5}$$

where $\phi_i = -(1-\lambda)$, $\theta_{0i} = \dfrac{\gamma_i}{1-\lambda_i}$, $\theta_{1i} = \dfrac{\delta_{10i}+\delta_{11i}}{1-\lambda_i}$, $\theta_{2i} = \dfrac{\delta_{20i}+\delta_{21i}}{1-\lambda_i}$, $\theta_{3i} = \dfrac{\delta_{30i}+\delta_{31i}}{1-\lambda_i}$,

$\theta_{4i} = \dfrac{\delta_{40i}+\delta_{41i}}{1-\lambda_i}$, $\theta_{5i} = \dfrac{\delta_{50i}+\delta_{51i}}{1-\lambda_i}$, and $\theta_{6i} = \dfrac{\delta_{60i}+\delta_{61i}}{1-\lambda_i}$.

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[9] The vast literature on economic growth suggests a large number of potentially important determinants. The control variables used herein are among the most robust of such determinants. With the PMG, estimated degrees of freedom problem soon arises unless the available time series is long enough. This explains the limit we have imposed on the number of explanatory variables considered.

where θ_{0i}, θ_{1i}, θ_{2i}, θ_{3i}, θ_{4i}, θ_{5i}, and θ_{6i}, measure the homogeneous long-run level effect of special interest group and the control variables, while δ_{11i}, δ_{21i}, δ_{31i}, δ_{41i}, δ_{51i}, and δ_{61i}, measure the heterogeneous short-run dynamics. As already noted, $\varphi_i < 0$ provides evidence that the long-run relationship exists. Thus, a significant and negative value of φ_i indicates co-integration between y_{it} and the independent variables.

IV. The Empirical Results

Although the autoregressive distributed lag model renders (panel) unit-root tests of the variables under study unnecessary as long as they are I(0) and I(1), we performed these tests nevertheless to ensure that no variable exceeds the I(1) order of integration and, thus, possibly resulting in inconsistent estimations (Asteriou and Monastiriotis (2004)). To do this, we applied two commonly used panel unit root tests. The first is by Im, Pesaran, and Shin (2003) (IPS, hereafter), and it is based on the assumption that all series are non-stationary under the null hypothesis but accounts for heterogeneity in the autoregressive coefficient, which is assumed to change freely among the states under study. The IPS provides two statistics: the $Z_{\text{tbar-IPS}}$ based on the hypothesis that the residuals are not serially correlated, and the $W_{\text{tbar-IPS}}$ used when the residuals are serially correlated.

The second panel unit root test applied is due to Hadri (2000). It is a variation of the well-known KPSS unit root test, and it has the null of no unit root in any of the cross-section of the units in the panel. Specifically, it uses individual OLS regressions of the dependent variable on a constant or a constant and a trend; gives test statistics that are distributed as standard normal under the null; and allows the error process to be homoskedstic or heteroskedastic across the cross-section units in the panel. Thus, it provides two Z-statistics, one derived from the Lagrange

Multiplier (LM) statistic for the homoscedasticity assumption, and the other using the LM statistic that is heteroskedasticity consistent.

Table 1 reports the results for both the IPS and Hadri tests. Panel A of the table shows that for both the $Z_{tbar\text{-}IPS}$ and $W_{tbar\text{-}IPS}$ statistics, the level values of all of the series are non-stationary; however, all the series are stationary at the 1% significance level of the first difference. Panel B of Table 2 also shows that the Hadri panel unit root test results confirm these conclusions. Thus, together, the two tests indicate that the variables under study are integrated of order one (I (1)).

Next, we determine whether a meaningful long-run relationship exists between the variables in our model. For this we adopt the commonly applied (Pedroni (1999, 2004) test, which accounts for heterogeneity by using specific parameters. This test offers seven panel statistics, six of which rejects the null of no co-integration when they have large negative values and one (the panel-v test) which rejects the null of co-integration when it has a large positive value. Panel A of Table 2 reports the results of this test first for the three variables of main interest in this study (economic growth (y), special interest groups (SIGs) activity, and the square of special interest groups (SISQ)), and panel B of the table shows the results for all of the variables included in our model (equation (5)). For the former, the results reject the null of no co-integration in each of the seven tests involved for each of the three variables of interest. This provides strong support that a co-integrating relationship exists between economic growth and special interest group activity.

Panel B of Table 2 reports the panel co-integration results when each of the eight variables in our model is used as a dependent variable. In these cases, the null of no co-

integration is rejected in either four or five of each of the seven tests involved. This supports the view that the variables under study do have a stable long-run relationship.

We now estimate our model (equation 5) with PMG, MG, and DFE and then apply the Hausman test to determine whether a significant difference exists between these estimators[10]. The null of this test is that the difference between PMG and MG or PMG and DFE is not significant. A probability value larger than 0.05 means a rejection of the null, making the PMG the preferred estimator as it is consistent and efficient.

The results of the PMG, MG, and the DFE estimators, along with those of the joint Hausman test are reported in Table 3. The results of this test indicate that the long-run homogeneity restriction imposed by the PMG is accepted at the 5 per cent significance level as their p-values are bigger than 0.05. This is hardly surprising though given that such restriction improves the precision of the reported PMG estimates over those of the MG. The PMG is, therefore, preferred over the MG and the DFE and, thus, will be the focus of the subsequent analysis even though we continue to refer to the MG and DFE results for comparison purposes.

Further evidence on the presence of a long-run relationship amongst the variables under study is shown by the coefficients of the error-correction (EC) terms in Table 3: -0.123, -0.425, and -0.221, for the PMG, MG, and DFE, respectively, all of which are negative, significant at the 1 per cent level, and are within the dynamically stable unit circle. Thus, the variables in our model are, indeed, co-integrated, giving credibility to our long run estimates[11]. Also, the error-correction terms provide some insights on how long it takes for an SIGs activity shock to go through its short-run path to eventually make its full long-run growth effects evident.

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[10] These and many of the related estimations were performed using the Stata 13 statistical software.
[11] The presence of a long-run relationship between growth and SIGs activity indicates that the two variables can be mutually causal. However, our ARDL specification (equation (5)) includes lagged explanatory variables, a feature which minimizes potential endogeneity or reverse causality problems (Huang, Fang, and Miller (2013)).

Specifically, the PMG suggests a relatively slow speed of adjustment of about 8 years while the MG and the DFE indicate a faster adjustment path of about 2.5 and 4.5 years, respectively.

Prior to analyzing the results for the main issues of interest herein we briefly describe the results relating to the control variables. The initial estimates of our model included the initial condition variable (initial income per capita). Although this variable has the expected negative sign it was not significant, and was dropped from the final estimations with no noticeable impact on the results. The lack of significance of this variable may be reflective of the observed large decrease in the rate of income convergence across the U.S. states over the past thirty years. One of the reasons suggested for this decrease is the large increase in housing prices and housing regulation in high-income areas and its negative impact on labor mobility (cf., Ganong and Shoag (2012)).

Also, the two human capital variables (HI and CO) were included together in the initial estimations, however, only CO was significant although each was positive and significant when entered alternatively. In the final estimates we omitted HI and retained CO, again with no discernible impact on the results. The human capital achievement variable (CO) indicates a robustly positive and significant long-run growth effect in all three estimators; however, it indicates a negative short-run impact which is marginally significant only in the case of the MG.

The tax burden variable has a significant negative long-run effect on growth in the PMG but it is not significant in the short run. The income inequality variable (GINI) indicates a strong positive long-run impact on growth in all three of the estimators. This is consistent with the findings of Partridge (1997 and 2005) for the U.S. states, as well as the results obtained by Knack (2003), Horgos and Zimmermann (2009), and Rossignoli (2013). However, the interactive term between income inequality and SIGs activity (GINI x SIGs) shows a negative

long-run effect in all three estimators though significantly so in only the PMG and the MG. This suggests that the detected positive effect of inequality on growth is lower for lower-income states.

Now we turn to the main issue of interest, namely, the relationship between SIGs activity and state economic growth. Table 3 shows that while the main results for the PMG and the MG are fairly similar, they are substantially different from those of the DFE. Specifically, while the PMG and the MG indicate an inverted U-shaped relationship, the DFE suggests a U-shaped relationship. For the PMG, the coefficients of SIGs activity and its square are 0.219 and -0.003 (0.311 and -0.007 for the MG), respectively, both of which are significant at the 1 percent level. These results are in accord with both the *diminishing marginal returns* and the *economic development views* (Choi (1983), Olson, Heckelman, and Coates (2003), and (Heckelman and Coates (2003)) which maintains that the marginal effects of SIGs activity are growth-enhancing in the early stages but become harmful to growth in the later stages.

A similar positive and then negative growth effect of SIGs activity is revealed by the PMG for the short run. However, the positive coefficient for SIGs is only marginally significant, while the negative coefficient of its square is significant at the 5 per cent level. But the economic significance of these short-run effects is small. Thus, overall, we can conclude on the basis of the PMG estimates that SIGs activity growth effects are more potent in the long-run than the short-run. Thus, our results are also consistent with *the farsighted view* alluded to above which maintains that special interest groups are mostly farsighted in their policy preferences because they attempt to maximize their members' expected welfare over the long run.

The pattern of both short- and long-run threshold SIGs activity growth effects revealed by the PMG in Table 3 are closely echoed by the MG results in the same table and, thus, suggesting robustness for the results. A tempting interpretation of this similarity in the results would be that even if the long-run coefficients differ across the states, as assumed by the MG, the revealed inverted U-shape relation between SIGs activity and growth would still be tenable. However this is not the case given that the Hausman test could not reject the long-run homogeneity assumption of the PMG.

The results for the DFE are reported in Table 4. In contrast to those for the PMG and the MG, these results suggest a long-run U-shaped relationship between growth and SIGs activity, as suggested by *the income distribution view*. Note, however, that the MG's SIGs negative coefficient is not significant but its SIGs square variable has a positive coefficient and it is significant at the 1% level.

As already stated, the quadratic method used herein to test for the presence of a nonlinear relation has recently drawn criticism from the important paper by Lind and Mehlum (2010). In the context of this paper, their criticism would be that even if the estimated coefficients of SIGs activity and its square are significantly positive and negative, respectively, in our estimated model, and we are also able to compute an extreme value that lies within the data range, these are only necessary but not sufficient conditions to prove that an inverted U-shaped relationship actually exists between SIGs activity and economic growth. The source of the problem is "….when the true relationship is convex but monotone. A quadratic approximation will then erroneously yield an extreme point and hence a U shape." (Lind and Mehlum (2007, p. 2)). To resolve the issue, Lind and Mehlum (2010) proposed a method that incorporates the exact necessary and sufficient conditions for the u-test of a U shape in both finite samples and for a

large class of models. We applied their u-test to determine whether in fact an inverted U-shape relationship exists between SIGs activity and economic growth. To do this, we estimated the following model (cf., Samargandi, Fidrmuc, and Gosh (2013)):

$y_i = aSIGs_i + bSIGs_i^2 + X_iC + error\ term$

To test the joint hypothesis:

$$H_0 : (a + b2SIGs_{min} \leq 0)\ U\ (a + b2SIGs_{max} \geq 0)$$

$$H_1 : (a + b2SIGs_{min} > 0)\ U\ (a + b2SIGs_{max} < 0)$$

where $SIGs_{min}$ and $SIGs_{max}$ are the minimum and maximum values of the SIG variable, respectively, and X_i are the control variables. The existence of a U shape is confirmed if the null hypothesis is rejected.

We implemented the test for a U-shaped relationship using the Stata 13 statistical software and the ado-file u-test provided by Lind and Mehlum. The results of the u-test indicate that the lower bound slope of SIG is positive (0.0070) while its upper bound slope is negative (-0.0124). More importantly, both are statistically significant with t-values of 2.66 and -3.45, respectively, indicating that the null hypothesis of no inverted U-shape is rejected. Also, the computed extreme value is given as 15.8, indicating that the threshold or turning point is within our sample period, and that this point is reached when the average SIGs activity (based on the percentage of each state's public and private nonagricultural wage and salary employees who are union members, and which ranges from 3.8% to 38.8%) is at the 15.8% level. Thus, the results of the u-test reaffirm our finding of an inverted U-shape relationship between special interest group activity and economic growth in the U.S. states over the 1975 to 2004 years.

V. Summary of results and conclusions

Previous empirical studies on the relationship between SIGs activity and economic growth have used cross-section regressions and different forms of cross-section time-series methods which can capture only the long-run growth effects, and which often assumes parameter homogeneity across the units being studied, possibly leading to biased parameter estimates (cf., Pesaran and Smith (1995)). This paper has extended the literature by separating and estimating the short-run and the long-run growth effects of SIGs activity. We do this by applying new methods for non-stationary panels that include panel unit root tests, a co-integration test, an appropriate U-shaped relationship test, and three suitable estimators (PMG, MG, and DFE) to a panel of data for 48 U.S. States over the 1975 – 2004 period. Both the co-integration test and the significantly negative error-correction coefficients of the PMG, MG, and DFE models suggest that a meaningful and stable long-run relationship exists between SIGs activity and economic growth. However, the Hausman test suggests that the PMG, which allows the long-run parameters on growth to be identical across the states but allows the short-run adjustments to long-run equilibrium to differ across the states, is preferred over the MG, which assumes that both the short-run dynamics and the long-run coefficients are heterogeneous across the states, and the DFE, which constraints all of the short-run and long-run coefficients to be the same across the states. Both the PMG and the MG estimators robustly reveal evidence for an inverted U-shape relationship between SIGs activity and growth. This is consistent with *the diminishing marginal returns and the economic development views* which suggest the existence of a threshold point below which special interest groups' lobbying and rent-seeking activities boost long-run growth performance but above which they have harmful effects on long-run growth

effort. Also, since the results suggest that the SIGs growth effects are more potent in the long run, they are in line with the farsighted view alluded to above.

Also of importance is that the results from the Lind and Mehlum (2010) test reaffirms the validity of the results obtained from the customary quadratic method used herein to test for nonlinearity. Additionally, their u-test suggests that the threshold point U.S. States face is reached when the proxy for SIGs activity (the percentage of each state's public and private nonagricultural wage and salary employees who are union members (which varies from 3.8% to 38.7%)) is at 15.8%.

To conclude, we state the obvious, namely, that data and methodological problems beset any attempt to analyze empirically the complicated SIGs activity – economic growth process. We have by no means resolved these problems. However, we believe that our preliminary attempt to disentangle the short- and the long-run effects of SIGs activity growth effects have provided some insights into the said process. With regard to future research, these areas, among others, warrant further investigation. First, we note that separating the short- from the long-run SIGs activity growth effects is worthwhile for obvious and important reasons. Much progress can be made in this area using various dynamic panel error correction models. Second, it is important to investigate whether alternative measures of special interest group activity will yield results that are similar or different to the ones obtained herein. Third, performing further robustness tests on our model specifications using data for the U.S. states and other spatial units may provide further insights on the issues. Finally, *the income distribution view* proposed in this paper, though not supported by the preferred model, seems intuitive and important in its own right and, thus, should be given more attention in future research.

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Table 1: Results of Panel Unit Root Tests

IM-Pesaran-Shin Test					
Panel A		$Z_{tbar-IPS}$		$W_{tbar-IPS}$	
Variables	Level	First Differences	Level	First Differences	Order of Integration
y	9.247	-20.445***	9.2034	-12.737***	I(1)
GINI	5.774	-18.872***	5.7439	-12.293***	I(1)
GINI*SIGs	-3.911	-27.955***	-2.7417	-15.426***	I(1)
SIGs	-1.030	-30.261***	-0.2368	-16.234***	I(1)
SISQ	-3.591	-28.396***	-3.2834	-15.281***	I(1)
SLTAX	-2.5624	-25.420***	-3.9263	-14.4237***	I(1)
HSCHOOL	-1.2456	-29.702***	-0.7936	-16.797***	I(1)
COLLEGE	7.050	-31.6***	7.0489	-16.724***	I(1)
Panel B		Hadri Test			
Variables	Level	Prob	First Differences	Prob	Order of Integration
y	123.67	0.0000***	1.409	0.0800	I(1)
GINI	121.49	0.0000***	19.18	0.0000***	I(1)
GINI*SIGs	70.51	0.0000***	-4.767	1.0000	I(1)
SIGs	109.01	0.0000***	-4.0791	0.9999	I(1)
SISQ	102.75	0.0000***	-3.2707	0.9994	I(1)
SLTAX	36.53	0.0000***	31.3736	0.0000***	I(1)
HSCHOOL	123.65	0.0000***	0.6962	0.2431	I(1)
COLLEGE	119.94	0.0000***	19.659	0.0000***	I(1)
*** indicates significance at the1% level.					

Table 2: Pedroni (1999, 2004) Panel Co-integration Test Results

Panel A							
Dependent variables	Panel V	Panel Rho	Panel PP	Panel ADF	Group Rho	Group PP	Group ADF
y	1.98**	-5.768***	-8.102***	-7.558**	-4.044***	-9.342***	-9.405***
SIGs	3.742***	-5.296***	-7.190***	-6.295***	-3.867***	-8.115***	-7.981***
SISQ	3.216**	-4.638***	-6.566***	-6.136***	-3.398**	-7.570***	-7.760***
Panel B							
Dependent variables	Panel V	Panel Rho	Panel PP	Panel ADF	Group Rho	Group PP	Group ADF
y	0.822	4.242	-3.191**	-3.966***	6.540	-2.868**	-4.053***
GINI	-1.720*	3.881	-4.049***	-3.763***	6.101	-3.831***	-3.246**
GINI*SIGs	-1.887**	3.225	-5.728***	-4.956***	5.369	-5.947***	-5.050***
SIGs	-0.9225	2.411	-8.348***	-6.835***	4.607	-8.629***	-7.159***
SISQ	-0.377	1.898	-11.45***	-9.180***	4.130	-12.28***	-10.19***
SLTAX	-1.862**	4.631	-2.623**	-2.535**	7.000	-2.026**	-1.391*
HSCHOOL	-0.9053	3.625	-6.568***	-6.0396***	5.819	-6.719***	-6.183***
COLLEGE	-1.716*	3.465	-6.597***	-6.522***	5.819	-7.068***	-7.112***

Note: The null hypothesis of the test is that there is no co-integration between the variables. The Pedroni (2004) tests have a critical value of -1.64 (k < -1.64 suggests rejection of the null), except for the v-statistic that has a critical value of 1.64 (k > 1.64) suggests rejection of the null). *, **, and *** indicates rejection of the null hypothesis at the 1%, 5%, and 10% levels of significance, respectively.

Table 3: Long-run and short-run effects of special interest groups on the economic growth of U.S. States

Dependent variable: log real income per capita ((Δlogy)	(1) Pooled MG	(2) MG	(3) DFE
Error-correction coefficient	-0.123** (0.019)	-0.425** (0.028)	-0.221** (0.015)
Long-run coefficients			
GINI	7.23** (0.670)	6.78** (1.739)	1.736** (0.339)
GINI x SIGs	-0.489** (0.058)	-0.529** (0.153)	-0.010 (0.019)
SIGs	0.219** (0.028)	0.311** (0.122)	-0.019 (0.011)
SIGSQ	-0.003** (0.0004)	-0.007* (0.003)	0.00036** (0.0001)
COLLEGE	1.011** (0.245)	1.290** (0.427)	1.571** (0.243)
SLTAX	-2.025** (0.533)	1.651 (1.160)	0.015 (0.013)
Short-run coefficients			
ΔGINI	-0.014 (0.280)	-0.787* (0.400)	0.046 (0.140)
ΔGINI x SIGs	0.395** (0.026)	0.074* (0.030)	0.018* (0.008)
ΔSIGs	-0.025* (0.013)	-0.038* (0.017)	-0.005 (0.003)
ΔSIGSQ	0.0004** (0.0001)	0.0006* (0.0002)	0.000009 (0.0003)
ΔCOLLEGE	-0.007 (0.061)	-0.226* (0.109)	-0.148 (0.086)
ΔSLTAX	0.176 (0.297)	-0.129 (0.313)	0.002 (0.002)
Joint Hausman Test		8.28 (0.2183)	2.03 (0.7300)
No. of observations	1440	1440	1440
No. of states	48	48	48
Time period	1975-2004	1975-2004	1975-2004
* and ** indicate significance at the 5% and 1% levels, respectively.			