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The econometrics of rational partisan theory

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This study develops an econometric intervention model representing the standard empirical approach to testing Alesina's (1987) Rational Partisan Theory implication that elections lead to short-term changes in output growth and unemployment. This intervention approach may be subject to two econometric difficulties. First, the cyclical nature of the autoregressive variables suggest the regression residuals may be serially correlated. Second, the election intervention variable may be endogenous to the cyclical variables. Empirical support for the model is mixed. Ordinary Least Squares estimates for both series produce a coefficient for the intervention variable which is of the predicted sign but not significant. The output growth regression results are robust to serial correlation and endogeneity concerns. For unemployment, controlling for serial correlation generates a significant coefficient, but adjusting for endogeneity does not.

I. INTRODUCTION

There has been a recent rekindling of interest in models of electoral cycles linking partisan policies to the state of the economy. The literature has grown so rapidly there are now several recent surveys, including Hibbs (1992), Gartner (1994a), Alesina (1995), Price (1997), and Paldam (1997), each of which focuses on a different aspect of the literature. One conclusion is clear from each of the authors: much of this revival is due to the impact of Alesina's (1987) Rational Partisan Theory (RPT) model which suggests election outcomes may trigger temporary changes in real macroeconomic variables. This model is based on political parties holding different policy goals which prevent rational agents from perfectly predicting prior to an election the subsequent monetary policy to be enacted since the election might result in a change in party power and consequently the inflationary target. Agents are instead assumed to forecast monetary policy as a weighted average of the parties' positions. A number of studies have attempted to test the central implication of conservative party victories leading to short-term recessions due to lower than expected inflation occurring, and liberal victories leading to temporary booms from higher than expected inflation. The empirical tests generally focus on real output growth and unemployment using intervention analysis on reduced form autoregressive (AR) functions, where the intervention variable denotes the victorious party in the election (Sheffrin, 1989; Alesina and Roubini, 1992; Gartner, 1994b).

Such a representation can lead to two primary econometric difficulties. First, it is well established in the time-series literature that regression analysis on a truncated AR representation requires the series to be integrated of a particular order, otherwise the innovations will be serially correlated making ordinary least squares (OLS) parameter estimates inefficient, or worse, inconsistent. The so-called 'unit-root hypothesis' literature has been unable to definitively conclude whether various macro series are trend-stationary (TS) or difference-stationary (DS). Second, OLS requires the absence of contemporaneous correlation between the explanatory variables and the regression residual. If the election outcome depends on current economic conditions the intervention variable will be correlated with the error term, resulting in inconsistent parameter estimates.

It is therefore necessary to consider both the order of integration and the possibility of endogenous elections when testing RPT. Using data from Britain on real GDP growth and unemployment, it is found that empirical representations estimated by OLS which follow the assumptions of stationarity and exogenous election out-

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comes yield only very weak support for the model, in the sense that the intervention variable is of the predicted sign but not close to being statistically significant. However, if unemployment is not stationary around a linear trend, so that the series is underdifferenced, then correction for serial correlation generates initial strong support for the model. However, in this case the Hausman test statistic rejects election outcomes as exogenous. Correcting for endogeneity with an instrumental variable (IV) estimator, the election outcome coefficient loses its significance and even reverses sign. The output growth equations do not reject the intervention variable as exogenous and the IV estimates of this parameter remain insignificant but of the predicted sign. This holds for both DS and TS representations of growth. Thus, the evidence on RPT, and endogenous election outcomes, is mixed. The results also suggest elections may depend on unemployment, but not on output growth.

The rest of the paper is structured as follows. In Section II, the RPT regressions for output growth and unemployment are derived. In Section III, the hazards of econometric estimation of the RPT model are developed. In Section IV, the empirical results are presented. Finally in Section V, the paper is concluded by summarizing the results.

II. THE RATIONAL EXPECTATIONS PARTISAN BUSINESS CYCLE MODEL

The regressions which motivate the econometric section of the paper are now generated. To do so, an RPT model is designed which directly results in an AR process for the economic indicators to coincide with empirical formulations typically used in the literature. This will make the discussion in Section III easier to outline.

First, consider the effect of an election on output growth, where the economy is modelled as follows.

$$\Delta y_t = \bar{y} + h(m_t - m_t^e) + \eta_t \qquad h > 0 \tag{1a}$$

$$\Phi(B)\eta_t = \nu_t \tag{1b}$$

$$\Phi(B) = 1 - \sum_{i=1}^{r} a_i B^i \tag{1c}$$

$$m_t = m^C C_t + m^L L_t \tag{1d}$$

$$m_t^e = m^C C_t^e + m^L L_t^e \tag{1e}$$

$$C_t^e = \begin{cases} \pi, & \text{election period} \\ C_t, & \text{else} \end{cases} \qquad L_t^e = \begin{cases} 1-\pi, & \text{election period} \\ L_t, & \text{else} \end{cases} \tag{1f}$$

In the above equations, y_t is log output; Δ is the first difference operator, such that $\Delta = 1 - B$, where B is the backshift operator such that $B^i x_t = x_{t-i}$; $\bar{y}$ is the full-employment growth rate; h is a parameter signalling the positive relationship between unanticipated monetary expansion and real growth; η_t and ν_t are correlated innovations; C_t and L_t are dummy variables for Conservative and Labour party-ruled governments such that $C_t = 1 - L_t$; m_t is the value of monetary policy instruments (e.g. monetary base growth); m^C and m^L are the optimal time-consistent monetary values for the two parties; π is the nonstochastic probability[1] that the Conservative party will win a majority of seats in a parliamentary election, where $0 < \pi < 1$; and parameters denoted by the superscript ^{e} are the expected values of the parameters.

Equation 1a represents an expectations-augmented Phillips Curve, where the innovation process is propagated through Equations 1b and 1c. Monetary policy is controlled by the governing party which adopts its preferred (fixed) policy targets in Equation 1d. When the money supply is unexpectedly large, output grows above trend. Similarly, an unexpectedly tight monetary policy results in a contraction of output growth.

Following the standard premise of Alesina's (1987) RPT, it is assumed agents base their wage and price contracts on the expected monetary policy. Thus, monetary policy is neutral in that only unanticipated changes in monetary policy alter the real economy in Equation 1a. The two major parties have different preferences on the control of monetary instruments (Equation 1d), where $m^C < m^L$ by assumption (see Hibbs (1992) for a summary of the evidence). Finally, expectations in Equations 1e and 1f are formed rationally by economic agents. They are aware of the differences in the parties' optimal monetary policy, and that elections will lead to a change in the monetary policy targets if the current government is replaced by the outside party.

In a nonelection period, it is known with certainty which party controls the government, thus $C_t^e = C_t$ and $L_t^e = L_t$ from Equation 1f. Therefore $m_t^e = m_t$, since Equations 1d and 1e are identical. Substituting Equations 1b–1f into Equation 1a yields:

$$\Delta y_t = \bar{y} + \Phi^{-1}(B)\nu_t \tag{2}$$

[1] It is assumed in Equation 1f that the election outcome cannot be perfectly forecast, possibly due to random preference shocks that create uncertainty as to the position of the median voter (Alesina and Rosenthal 1995). The assumption of a nonstochastic probability has been criticized by Chappell and Keech (1988) since this leads to a fixed 'surprise' outcome in each election. A stochastic element will be introduced by the instruments for the election outcome variable as explained in Section III.

and rearranging terms,

$$\Phi(B)\Delta y_t = \Phi(B)\bar{y} + \nu_t \tag{3}$$

Expanding Equation 3 and rearranging terms yields,

$$\Delta y_t = k + \sum_{i=1}^{r} a_i \Delta y_{t-i} + \nu_t \tag{4}$$

where $k = \bar{y}(1 - \sum a_i)$. Equation 4 models output growth as a standard AR(r) distribution.

The stochastic process changes in an election period since it is not known which party will control monetary policy in the future. To minimize their error when setting contracts, rational economic agents form weighted expectations of the two parties' preferred monetary policy target as given in Equations 1f and 1e. The impact of partisan monetary policies, given by $h(m_t - m_t^e)$ in Equation 1a, can be summarized as follows. If the Labour party emerges victorious from the election, $h(m_t - m_t^e) = h\pi(m^L - m^C) > 0$, whereas when the Conservative party is victorious, $h(m_t - m_t^e) = h(1-\pi)(m^C - m^L) < 0$. Thus, actual monetary policy will be more expansionary than expected when the Labour party wins and less expansionary than anticipated when the Conservative party wins. The model thus predicts electoral cycles of higher output growth immediately following a Labour victory and vice versa for a Conservative victory. In the periods in between elections, $h(m_t - m_t^e) = 0$. Unlike the traditional Partisan Theory model pioneered by Hibbs (1977), the effect on growth is expected to be short-lived as new contracts formed after the election will be adjusted to accurately reflect the current party in power.

This election effect can be represented using Box–Tiao intervention analysis on Equation 4 which results in an empirical formulation coinciding with Alesina and Roubini's (1992) time-series representations. In an election period, the stochastic process of output growth is altered in a manner dependent upon the election outcome. Under the assumption that $\pi = 1/2$, the election surprise will yield a symmetric effect on the economy, which can be represented as

$$\Delta y_t = k + \sum_{i=1}^{r} a_i \Delta y_{t-i} + \lambda E_t + \nu_t$$

$$E_t = \begin{cases} +1. & \text{Labour victory} \\ -1, & \text{Conservative victory} \\ 0, & \text{no election} \end{cases} \tag{5}$$

where $\lambda = h(m^L - m^C)/2$. In this representation of the election outcome intervention variable, E_t, RPT predicts $\lambda > 0$. Equation 5 represents a typical simple test of RPT, although the representation of E_t differs slightly. Alesina (1989) and Alesina and Roubini (1992) coded their intervention variable similarly, but only for actual changes in government party control and thus their model was not a direct test of RPT. As Sheffrin (1989) notes, that type of variable representation is valid only if there is no uncertainty when the incumbent party is reelected. Hibbs (1992) argues further that their election variable formulation is empirically indistinguishable from an adaptive expectations model.

A similar relationship to Equation 5 can also be derived to consider the effect of an election on unemployment.

$$\begin{aligned} u_t &= \bar{u}_t + f(m_t - m_t^e) + \xi_t, \\ \bar{u}_t &= \bar{u} + \delta t \qquad \delta \geqslant 0, \qquad f < 0 \end{aligned} \tag{6a}$$

$$\varphi(B)\xi_t = w_t \tag{6b}$$

$$\varphi(B) = 1 - \sum_{i=1}^{q} b_i B^i \tag{6c}$$

Here, u_t represents the unemployment rate; $\bar{u}_t$ is the stochastic natural rate of unemployment; $\bar{u}$ is the initial rate of unemployment; δ is a small non-negative value allowing the natural rate of unemployment to slowly increase over time; f is a parameter denoting an inverse relationship exists between unemployment and unanticipated money growth; and ξ_t, w_t are innovations similar to those in Equations 1a and 1b.

Substituting the policy Equation 1d and the expectations functions (Equations 1e and 1f), along with Equations 6b and 6c, into Equation 6a and following the same analysis as above for output growth yields the Box–Tiao intervention model

$$u_t = \alpha + \tau t + \sum_{i=1}^{q} b_i u_{t-i} + \gamma E_t + w_t \tag{7}$$

where $\alpha = \bar{u}(1 - \sum b_i) + \delta \sum i b_i$, $\tau = \delta(1 - \sum b_i)$, and $\gamma = f(m^L - m^C)/2$. RPT predicts $\gamma < 0$ since higher output growth from unanticipated money growth should be accompanied by reduced unemployment ($f < 0$).

Equations 5 and 7 represent the basic regressions for the tests of RPT. But as will be shown in the next section, there are some potentially serious econometric problems the RPT literature has traditionally overlooked.

III. ECONOMETRIC CONSIDERATIONS

There are two primary concerns which hamper traditional OLS estimation of Equations 5 and 7. First, the output growth and unemployment series must be $I(0)$ for OLS to be consistent and efficient. Second, the explanatory variables must not be correlated with the error term. The AR portion of the regressions (constant and lagged dependent variables) are by definition not contemporaneously correlated with the regression residual, but that claim cannot be made for the election outcome, the key variable under consideration here. We now consider both these problems and offer solutions.

The order of integration

Under the assumption that $y_t \sim I(1)$, we can expect $\nu_t \sim (0, \sigma_\nu^2)$. If one were only interested in forecasting output growth and did not consider the potential election effects, Equation 3 would seem a proper model, but only if y_t were in fact DS. If however, y_t is TS, its representation is:

$$y_t = \kappa + \theta t + \sum_{i=1}^{r} a_i y_{t-i} + \varepsilon_t \tag{8}$$

To find the proper analogue to Equation 4 note,

$$y_{t-1} = \kappa + \theta(t-1) + \sum_{i=1}^{r} a_i y_{t-i-1} + \varepsilon_{t-1} \tag{9}$$

and thus

$$\Delta y_t = \theta \Delta t + \sum_{i=1}^{r} a_i \Delta y_{t-i} + \Delta\varepsilon_t \tag{10}$$

which implies

$$\Delta y_t = k + \sum_{i=1}^{r} a_i \Delta y_{t-i} + \nu_t \tag{11}$$

where $k = \theta$ and $\nu_t = \varepsilon_t - \varepsilon_{t-1}$.

Computationally, Equation 11 is indistinguishable from Equation 4. This makes Equation 4 a consistent representation for output growth when log output is either TS or DS. But it should be clear from Equations 10 and 11 that ν_t is a unit root moving average (MA) process when y_t is TS. In this case, OLS on Equation 4 (or equivalently, Equation 11) yields inefficient parameter estimates.

Similarly, OLS on Equation 7 is also only valid when unemployment is a stationary (or trend-reverting) series. If $u_t \sim I(1)$ then it would be necessary to first difference the unemployment series to achieve stationarity. Following the procedure above, and again temporarily ignoring the potential election outcome effect,

$$\Delta u_t = \tau + \sum_{i=1}^{q} b_i \Delta u_{t-i} + e_t \tag{12}$$

where $e_t = w_t - w_{t-1}$.

If $u_t \sim I(1)$, then e_t is a standard white noise error process. In this case, since $w_t = w_{t-1} + e_t$, recursive substitution can show $w_t = \sum_{i=0}^{\infty} e_{t-i}$ which means the error process in Equation 7 is nonstationary. Although w_t is still a zero mean process, its variance is undefined and thus the sampling distribution on the coefficient estimates does not have finite second moments.

Table 1. *Econometric considerations for overdifferencing and underdifferencing*

	True process	Estimation in Levels	Estimation in Differences
	$I(0)$		
Coefficient estimates		consistent, efficient	consistent, inefficient
Regression residuals		white noise	serially correlated MA(1) unit root
	$I(1)$		
Coefficient estimates		inconsistent	consistent, efficient
Regression residuals		serially correlated nonstationary AR(1)	white noise

The concern of over- versus under-differencing the data can be summarized as follows. Using generic notation, if the true model is

$$Y_t = X_t\beta + \zeta_t \tag{13}$$

such that $\zeta_t \sim N_{\mathrm{ID}}(0, \sigma_\zeta^2)$, but a regression is run on the false model

$$\Delta Y_t = \Delta X_t\beta + \psi_t \tag{14}$$

then Equation 14 is over-differenced since $\psi_t = \zeta_t - \zeta_{t-1} \sim$ MA(1) unit-root process. OLS estimation will not be efficient.

Conversely, if the true model is Equation 14 such that $\psi_t \sim N_{\mathrm{ID}}(0, \sigma_\psi^2)$, but a regression is run on the under-differenced series Equation 13, then $\zeta_t = \sum_{i=0}^{\infty} \psi_{t-i} = (1-B)^{-1}\psi_t \sim \mathrm{MA}(\infty)$, which is asymptotically equivalent to an AR(1) process (Blough, 1992). OLS estimation in this case will not be consistent.

The pitfalls of OLS estimation are summarized in Table 1. Unless the researcher is confident of the order of integration, OLS may not be an appropriate estimator. Unit-root tests are often implemented to detect the presence of a nonstationary series. However, the power of unit-root tests are notoriously low. McCallum (1993) notes that unit-root tests have low power against a TS series with slow trend reversion.[2] Rudebusch (1993) finds that unit-root tests are also unable to distinguish between TS with low persistence, and DS models. Using augmented Dickey–Fuller tests, he is unable to reject the null of DS, but also cannot reject the null of TS. Thus, he claims the tests are

[2] Dickey *et al.* (1986) also express concern about the low power of Dickey–Fuller tests if the process is non-zero mean or a trend is included.

inconclusive. Harvey (1993, pp. 133–4) shows that augmented Dickey–Fuller tests are biased if an MA term is present since the underlying process is not a finite AR, and discounts various modifications which have been proposed for this case.

No general conclusion arises from the vast unit-root hypothesis literature regarding methodology or empirical implications. In fact, Rudebusch (1993, p. 27) claims, 'a new consensus should be formed that stresses the uncertainty about the existence of a unit root in real output and the uncertainty about the persistence of macroeconomic shocks.' Many others also claim it is not possible to tell with the existing data.[3]

Because of the controversy which continues to surround the unit-root hypothesis, some authors have advocated differencing each series as a precaution (Nelson and Kang, 1984; Dickey *et al.*, 1986) since the problems associated with over-differencing are considered less severe than for under-differencing, or running regressions in both levels and differences and then comparing the results (Hamilton, 1983; Dickey *et al.*, 1986). While this approach seems reasonable for studying the persistence of random innovations in the series, it is not a viable option for the RPT intervention model. According to RPT, elections have predictable effects on output growth and unemployment which do not persist. This claim does not hold for the level of output or for changes in unemployment. For example, as noted by Gartner (1994b), transitory changes to output growth following elections imply permanent changes in the level of output. Therefore, adding the intervention variable E_t to Equations 8 or 12 would not be a proper test of the RPT model. The intervention variable properly belongs in a truncated AR on output growth and unemployment but there is no guarantee these series contain stationary error processes.[4,5]

McCallum (1993) investigates the problem of over- and under-differencing and concludes the problems are serious for OLS but can be eliminated when properly accounting for the serial correlation of the error structure. Citing various Monte Carlo studies, he concludes 'that neither overdifferencing nor underdifferencing leads to serious estimation or testing mistakes in regression models with exogenous regressors, provided that the investigator takes intelligent account of serial correlation present in the regression residuals' (p. 30). His own numerical analysis supports this conclusion. He finds that OLS on an underdifferenced or overdifferenced series leads to parameter estimators which vary greatly from the true model, but AR(1) correction to the underdifferenced series and MA(1) correction to the over-differenced series yield almost identical results to OLS estimates on the stationary series. Thus, 'the presence or absence of differencing is not crucial when serial corrections are applied' (p. 31). McCallum's analysis suggests an alternative approach to determining the order of integration is through the use of serial correlation tests. This procedure is hindered under RPT since, as will be argued in the next section, there may also be contemporaneous correlation between the error term and the intervention variable, in which case the tests do not retain their asymptotic properties.

McCallum's results suggest we can still apply the RPT intervention to Equations 5 and 7. Since it is not clear *a priori* if $y_t \sim I(1)$ or $u_t \sim I(0)$ as assumed by these models, it is necessary to estimate Equation 5 using OLS which is valid under the null of DS, and also using an MA(1) unit-root correction which is valid under the alternative of TS. Similarly, both OLS and an AR(1) correction are used on Equation 7 which holds when unemployment is TS or DS, respectively. In each case only one of the regressions is valid. Thus, we can be confident of the results for each series only when both estimators yield the same conclusions. Here, we are concerned primarily with the sign and significance of the estimated parameter on the intervention variable E_t.

Endogenous elections

There is a secondary problem associated with regressions on Equations 5 and 7. Regardless of any correction for serial correlation, the estimated coefficients are only consistent if E_t is uncorrelated with the regression residual. Recent studies on RPT suggests this may not be the case. The discussion which follows will concentrate on the unem-

[3] See the references in McCallum (1993). Perron (1989) claims a consensus exists that unemployment is stationary, but the hysteresis literature (Lindbeck 1993; articles in Cross 1988) and some of the references above would not support this conclusion. There is also a symposium issue in the *Journal of Economic Perspectives* (Winter 1997) devoted to the topic of a natural rate of unemployment. Most of these studies do not support the notion of stationary long-term unemployment rates.

[4] In addition the Hausman test procedure, which will be used to determine if E_t is an exogenous variable, is only valid in the absence of serial correlation. Although inefficient estimators caused by the MA(1) process are in general less serious than the possibly inconsistent estimators created from the MA(∞) (or AR(1)) process in an under-differenced series, they are problematic for the Hausman procedure.

[5] Alesina and Roubini (1992) use a modified version of Equation 7 to ensure stationarity of their unemployment series. They transform their dependent variable as the deviation from the average OECD unemployment rate. However, RPT does not directly address this variable since the average OECD unemployment rate will presumably depend on the presence of elections in the other nations as well. For example, Conservative victories in other nations will increase the average unemployment rate, and thus British unemployment may fall below the average value even if there is no election in Britain. Elections in other nations will affect the stochastic process of the transformed unemployment series for any given country. A single intervention variable will not properly capture this effect.

ployment equations but an analogous rationale can also be provided for the output growth model.

The representation of unemployment in Equation 7 suggests unemployment is subject to cyclical variation in non-election periods as part of the normal business cycle. Balke (1991) modifies the standard RPT approach by introducing rational voters who realize the new electoral cyclical effects caused by each party upon (re)gaining control of policy as outlined in Section II and therefore support the party which will stabilize the economy. Thus, the winning party is expected to be a function of current economic conditions at the time of the election. Empirical work by Swank (1995, 1998) shows that in US elections, the Democratic party benefits from high unemployment and Republicans benefit when inflation is high. Sheffrin (1989) claims RPT models are not subject to the same endogeneity problem which plagues traditional opportunistic political business cycle models where the timing of an election in parliamentary-style nations may be manipulated by incumbents to secure reelection during periods of high growth.[6] While this may be true, the election calendar may be manipulated by partisan concerns. The implication from Balke (1991) on this point is that the Conservatives would be more likely to call an election in a boom period when inflation is considered high and the Labour party would prefer to hold an election during a recession when unemployment is considered to be relatively high. Thus the timing of the election would depend on the economic environment as well as which party controls the Prime Ministership, and the new composition of parliament following the election would depend on the economy.

The estimation problem can be shown as follows.[7] To simplify the notation, let

$$u_t = \alpha + Q_t B + w_t \tag{15}$$

where $Q_t = (t, u_{t-1}, \ldots, u_{t-q}, E_t)'$ and $B = (\tau, b_1, \ldots, b_q, \gamma)$ from Equation 7. The estimate of the B vector may be biased under OLS estimation since (dropping time subscripts for ease of notation)

$$\mathrm{E}(B_{OLS}) - B = \mathrm{E}(Q'Q)^{-1}Q'u - B = (Q'Q)^{-1}\mathrm{E}(Q'w)$$

From the discussion above, it is possible there is contemporaneous correlation between the election variable and the regression residual, i.e. $\mathrm{E}(E_t w_t) \neq 0$.

Now define Z_t to be a vector of instruments. The bias from IV estimation is

$$\mathrm{E}(B_{IV}) - B = \mathrm{E}(Q'Z(Z'Z)^{-1}Z'Q)^{-1}Q'Z(Z'Z)^{-1}Z'u - B$$
$$= (Z'Q)^{-1}\mathrm{E}(Z'w)$$

Table 2. *Instrument relevance for election outcome*

	ρ	R^2	$T \times R^2$	$1/\hat{\rho}^2$
Conservative	−0.062	0.011	1.53	89.29
Labour	0.139	0.026	3.64	37.61
Count	−0.266	0.078	10.72	12.78
CONSERVATIVE × *COUNT*	−0.272	0.081	11.14	12.30
LABOUR × *COUNT*	−0.199	0.047	6.42	21.34

Notes: Instrument definitions are in the text. ρ is correlation with Election outcome variable. R^2 is from bivariate regression on Election outcome variable. $\hat{\rho}^2 = [(T-1)/(T-2)]R^2 - 2/(T-2)$ is a relevance measure from Nelson and Startz (1990).

Thus, for the right set of instruments, IV estimation can potentially yield unbiased estimators. The trick, of course, is to find a set of instruments such that $\mathrm{E}(Z_t' w_t) = 0$ and where the extra instruments in Z_t do not have any causal influence on u_t (i.e., they are restricted to have coefficients equal to zero if included in Equation 15).

Here the poll support for each party and the length of time since the last election are considered as possible instruments. Poll data is a good, although imperfect, predictor of election outcomes (Chappell, 1990). Since there is always a nonzero probability of losing power in any election, elections may be temporarily postponed by the incumbent party until the marginal cost of holding an election equals the marginal benefit from a new term to restart the electoral term clock (Balke, 1990).

In the simplest one variable-one instrument case, Nelson and Startz (1990) suggest a suitable instrument exists when $TR^2 > 2$ from the first-stage regression. In addition, they recommend testing for $1/\hat{\rho}^2 < T$, where $\hat{\rho}^2 = [(T-1)/(T-2)]R^2 - 2/(T-2)$. Bound *et al.* (1995) and Shea (1997) base their relevance criteria on the partial R^2 measure.

Tests for instrument relevance are presented in Table 2. Monthly Gallup poll data for Conservative and Labour party preference rate in the UK through June, 1994 is available from Butler and Butler (1994). To match the economic data from OECD, the sample begins in 1960. The values are aggregated to a quarterly rate (only April 1966 and July 1970 are missing values). The election variable is constructed from the appendix in Alesina and Roubini (1992) and updated from Banks (1996). Following the analysis in Alesina and Sachs (1988) and Alesina (1995), where it is shown that RPT effects in the USA last for two years following each election, the election intervention variable retains its +1 or −1 value for eight quarters.[8] Not surprisingly, the party popularity variables are not

[6] For empirical evidence supporting endogenous election timing in Britain and Japan, see Heckelman and Berument (1998).

[7] An excellent overview of simultaneity as applied to time-series models is provided by Hamilton (1994, chapter 9).

[8] The only exception is for the 1974:1 election when a new election followed in 1974:3. Alesina (1989) also uses an eight quarter specification. Alesina and Roubini (1992) reports results using a six quarter specification but note that four or eight quarter specifications yield similar results.

strongly correlated with the election variable since party popularity fluctuates in nonelection periods when the election variable always takes the value of zero.[9] They are, however, of the expected sign. The negative correlation of *CONSERVATIVE* and the positive correlation of *LABOUR* suggests the Labour party is less likely to emerge victorious following an election (+1 for the election variable) the greater the Conservative popularity and more likely to be victorious the greater the Labour party citizen support. As instruments, they each individually pass the relevance tests ($T = 138$). The *COUNT* variable, which measures the number of quarters since the last election, is more strongly correlated with the intervention variable since it better captures the timing an election (smaller values of *COUNT* tend to occur when $E_t = 0$) but not the sign of the variable in an election period.

To generate as strong an instrument as possible from this list, interactions between *COUNT* and the party support ratings are next considered. The interaction with *CONSERVATIVE* improved the degree of correlation and produced the lowest $1/\hat{\rho}^2$ value but the interaction with *LABOUR* yields a weaker correlation compared to *COUNT* by itself. Thus the best instrument from this list is taken to be *CONSERVATIVE* × *COUNT*, but the R^2 suggests this is still a 'weak' instrument (Nelson and Startz, 1990). As suggested by Bound *et al.* (1995) we note the increase in R^2 from adding the weak instrument to the first-stage regression for the output growth representation is 0.06 (t-ratio $= -2.78$) and for the unemployment representation is 0.08 (t-ratio $= -3.43$). These values are the highest of all the instrument choices and suggest along with the relevance tests in Table 2 that the instrument, although 'weak', may still be suitable.

A final consideration might be to include all the potential instruments since they each pass the instrument relevance tests. The test for overidentified restrictions, where the null is all the extra instruments are exogenous, is analogous to a Lagrange Multiplier test where $TR^2 \sim \chi^2(m)$ in a regression of the second-stage residuals and m is the number of overidentifying restrictions (Greene, 1993). The overidentification statistic easily falls into the bounds of nonrejection for each specification. But adding additional instruments is not always the optimal procedure. Buse (1992) proves that more instruments need not lead to a reduced bias and may inflate the standard errors in IV estimation, and Hamilton (1994) proves that the IV estimator is the same as the Full Information Maximum Likelihood estimator only when the model is just identified. Furthermore, adding the additional instruments does not yield a discernable improvement to the partial R^2 measure when adjusted for degrees of freedom.[10] Since the conclusions regarding RPT are not altered between the just identified and over identified IV estimates, to conserve space only the IV regressions from the just identified model where *CONSERVATIVE*× *COUNT* is the only extra instrument are presented below.

IV. EMPIRICAL ESTIMATION OF THE MODEL

We now turn to the tests for RPT, concentrating here on the sign and significance of the election variable coefficient. According to the model, the coefficient on E_t should be positive for the output growth equations and negative for the unemployment equations. For both variables the stationarity and endogeneity issues are considered. Tests on RPT in Britain for output growth are presented in Table 3 and for unemployment in Table 4. Real GDP and unemployment rates for the civilian population are taken from the OECD database accompanying RATS 4.1. Real GDP is converted to annual growth rates for each quarter. The unemployment rate is converted to the quarterly average of each three month period. The number of lags is determined by the Schwarz Bayesian Information Criteria, which suggests output growth is AR(6) and unemployment is AR(3). In Tables 3 and 4, Columns (1) and (2) assume the election variable is exogenous and columns (3) and (4) assume E_t is endogenous. The odd numbered columns assume real GDP growth and unemployment are $I(0)$ so that Equations 5 and 7 are properly differenced. The even numbered columns correct for the possibility of overdifferencing output and underdifferencing unemployment.

It is well established that the Durbin–Watson statistic is biased against rejecting the null of no serial correlation when there are lagged values of the dependent variable included in the regression. Durbin's h-statistic, which is often employed in this case, suffers from being undefined when the estimated variance of the OLS estimate of the lagged dependent variable is greater than the inverse of the number of observations. Instead, consider the regressions

$$\begin{aligned} \nu_t = {} & \text{constant} + c_{y1}\Delta y_{t-1} + c_{y2}\Delta y_{t-2} + c_{y3}\Delta y_{t-3} \\ & + c_{y4}\Delta y_{t-4} + c_{y5}\Delta y_{t-5} + c_{y6}\Delta y_{t-6} + c_{y7}E_t \\ & + c_{y8}\nu_{t-1} + \text{error} \end{aligned} \tag{16}$$

$$\begin{aligned} w_t = {} & \text{constant} + c_{u0}t + c_{u1}u_{t-1} + c_{u2}u_{t-2} + c_{u3}u_{t-3} \\ & + c_{u4}E_t + c_{u5}w_{t-1} + \text{error} \end{aligned} \tag{17}$$

Two equivalent representations of Durbin's m-test is to either consider the statistical significance of c_{y8} and c_{u5}, or

[9] The results for *CONSERVATIVE* differ from *LABOUR* since although the Conservative and Labour parties are the two dominant political parties in Britain, the Liberal party garners enough support to prevent these two party preference ratings from summing to one.
[10] Shea (1997) suggests utilizing a correction of partial R^2 for degrees of freedom of the form $1[(T-1)/(T-n)]^*(1-R^2)$ where T is the number of observations and n is the number of instruments.

Table 3. *Tests for RPT on real GDP growth under alternative specifications of error term properties using* $\Delta y_t = k + a_1\Delta y_{t-1} + a_2\Delta y_{t-2} + a_3\Delta y_{t-3} + a_4\Delta y_{t-4} + a_5\Delta y_{t-5} + a_6\Delta y_{t-6} + \lambda E_t + \nu_t$

	(1) $\mathrm{Corr}(E_t, \nu_t) = 0$ $\nu_t \sim (0, \sigma_\nu^2)$	(2) $\mathrm{Corr}(E_t, \nu_t) = 0$ $\nu_t \sim \mathrm{MA}(1)$	(3) $\mathrm{Corr}(E_t, \nu_t) \neq 0$ $\nu_t \sim (0, \sigma_\nu^2)$	(4) $\mathrm{Corr}(E_t, \nu_t) \neq 0$ $\nu_t \sim \mathrm{MA}(1)$
k	0.271	0.271	0.274	0.274
	(3.56)	(5.97)	(2.31)	(2.62)
Δy_{t-1}	1.51	1.51	1.51	1.51
	(17.49)	(20.16)	(15.14)	(14.35)
Δy_{t-2}	−0.631	−0.631	−0.628	−0.628
	(−4.27)	(−3.96)	(−3.61)	(−3.09)
Δy_{t-3}	0.250	0.250	0.249	0.249
	(1.71)	(1.47)	(1.69)	(1.46)
Δy_{t-4}	−0.664	−0.664	−0.663	−0.663
	(−4.54)	(−3.39)	(−4.39)	(−3.34)
Δy_{t-5}	0.738	0.738	0.737	0.737
	(5.06)	(2.79)	(4.83)	(2.71)
Δy_{t-6}	−0.332	−0.332	−0.333	−0.333
	(−3.92)	(−2.37)	(−3.71)	(−2.34)
E_t	0.0533	0.0533	0.0661	0.0661
	(0.816)	(1.08)	(0.182)	(0.198)
σ_ν	0.478	0.478	0.478	0.478
SSR	27.44	27.44	27.45	27.45
Durbin's *m*				
t-ratio	−1.75		−1.52	
LM	3.22		3.57	
H-stat	0.00127	0.00839		

Notes. *t*-ratios listed in parentheses. E_t is an intervention variable with values of +1 for Labour party victory, −1 for Conservative party victory, and 0 for no election. The instrument for E_t in (3) and (4) is the Conservative party support rating interacted with the number of periods since the preceding election. MA(1) representations in (2) and (4) assume a unit root in the moving average process. LM is distributed as $\chi^2(1)$ with critical value = 3.84 at 0.05 level and 2.71 at 0.10 level. *H*-stat is the Hausman test statistic distributed as $\chi^2(8)$ with critical value = 15.51 at 0.05 level and 13.36 at 0.10 level for test of exogeneity of E_t.

calculate a modified LM test which under the null of no serial correlation is $TR^2 \sim \chi^2(1)$. The possibility of contemporaneous correlation with the election variable hampers these tests, which are only valid if the only misspecification involves serial correlation. It is not clear these tests retain their asymptotic properties under IV estimation.

With these caveats in mind, the LM version of the test rejects the null of no serial correlation (at the 10% level) for output growth, but the *t*-ratio test only corroborates this when elections are assumed to be exogenous (Column (1) in Table 3). There is no supporting evidence for the notion of serial correlation for the unemployment series. Under the standard RPT tests assuming exogeneity of elections and stable series for output growth and unemployment, the election intervention variable has the predicted sign, but is not significant. This corresponds to Alesina and Roubini's (1992) finding for the UK. Although their pooled regressions generated statistically significant coefficients (as mentioned above only considering changes in government rather than each election), the individual country regressions typically did not.

The presence of serial correlation for output suggests it may be overdifferenced, in which case the MA(1) correction will make the estimates more efficient. However, even in this case, the *t*-ratio for the election variable is only slightly improved, and still does not come close to statistical significance at conventional levels.

If, despite the serial correlation test results, unemployment is not TS, then an AR(1) correction is needed. If elections are exogenous (Column (2) in Table 4), then the intervention variable retains its negative sign and is significant. This lends more support to RPT, but the 'temporary' election effect will persist since unemployment does not revert to trend after temporary shocks. However, the Hausman test statistic is strongly significant in this case, suggesting the intervention variable is endogenous to the unemployment rate and therefore the coefficient is still not consistent. The instrumental variable estimate for this representation (Column (4)) does not generate a statistically significant coefficient and in fact reverses the sign.

The Hausman statistic tests for the exogeneity of E_t by comparing the IV estimator to estimates if E_t is exogenous.

Table 4. *Tests for RPT on unemployment under alternative specifications of error term properties using* $u_t = \alpha + \tau t + b_1 u_{t-1} + b_2 u_{t-2} + b_3 u_{t-3} + \gamma E_t + w_t$

	(1) $\text{Corr}(E_t, w_t) = 0$ $w_t \sim (0, \sigma_w^2)$	(2) $\text{Corr}(E_t, w_t) = 0$ $w_t \sim \text{AR}(1)$	(3) $\text{Corr}(E_t, w_t) \neq 0$ $w_t \sim (0, \sigma_w^2)$	(4) $\text{Corr}(E_t, w_t) \neq 0$ $w_t \sim \text{AR}(1)$
α	0.00661 (0.696)	0.0403 (0.439)	0.0118 (0.423)	0.00956 (0.194)
t	0.000571 (2.37)	0.00406 (2.51)	0.000578 (2.31)	0.00227 (2.35)
u_{t-1}	2.46 (2.37)	1.51 (17.41)	2.47 (31.37)	1.67 (17.64)
u_{t-2}	−2.02 (−13.94)	−0.416 (−2.69)	−2.02 (−13.51)	−0.626 (−3.38)
u_{t-3}	0.545 (7.38)	−0.166 (−1.87)	0.524 (7.02)	−0.0741 (−0.679)
E_t	−0.00481 (−0.655)	−0.0248 (−1.96)	−0.0243 (−0.246)	0.0171 (0.338)
σ_w	0.0518	0.0511	0.0532	0.0535
SSR	0.348	0.331	0.367	0.358
Durbin's m				
t-ratio	−1.28		−0.938	
LM	1.75		1.45	
H-stat	0.0390	18.89		

Notes. t-ratios listed in parentheses. E_t is an intervention variable with values of +1 for Labour party victory, −1 for Conservative party victory, and 0 for no election. The instrument for E_t in columns (3) and (4) is the Conservative party support rating interacted with the number of periods since the preceding election. AR(1) correction in (2) and (4) uses Cochrane–Orcutt method. LM is distributed as $\chi^2(1)$ with critical value = 3.84 at 0.05 level and 2.71 at 0.10 level. H-stat is Hausman test statistic distributed as $\chi^2(6)$ with critical value = 12.59 at 0.05 level and 10.64 at 0.10 level for test of exogeneity of E_t.

Under the null, the difference between the two sets of estimators should be small since the least-squares estimators would be consistent. The calculated Hausman statistics suggest the null of exogeneity cannot be rejected at any reasonable level except for the case of DS unemployment. Of course, the Hausman tests as applied here are only valid if the assumptions concerning serial correlation of the residuals are correct in each specification.

Thus RPT is only supported if unemployment is in fact DS and elections are exogenous, and not at all for output. This result emphasizes the impact of partisan cycles since although the uncertainty regarding partisan control of macroeconomic policy is only temporary, unemployment does not revert back to trend. If unemployment is TS, then temporary shocks to unemployment would only have a temporary effect as truly predicted by RPT, but in this case there is less empirical support for an election effect on unemployment.

V. CONCLUSIONS

In this paper, an econometric representation has formally been derived which coincides with many of the standard time-series applications of RPT tests. The analysis suggests these formulations may suffer from residuals which are serially correlated, and/or contemporaneously correlated with the election intervention variable. Previous empirical studies on RPT have tended to ignore these considerations.

Standard OLS analysis on British GDP growth and unemployment rates using an AR distribution only weakly supports the RPT model. Correcting for serial correlation does not directly alter the conclusions here. For output, the RPT variable retains its predicted sign but is never statistically significant regardless of the specification. For unemployment, a serial correlation adjustment generates a statistically significant coefficient but a Hausman test suggests the RPT variable is endogenous in this case. A further correction for endogeneity using an IV estimator generates an RPT coefficient that is not significant, and even reversed sign.

Although only one country is directly tested here, the implications are similar for other countries. To conclude, it is suggested that RPT tests of other democratic nations need to be undertaken again, using econometric estimation techniques that will generate consistent estimates in the presence of serially correlated innovations and endogenous election timing and outcomes, before conclusions regarding the viability of RPT are drawn.

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