General elections and government expenditure cycles:
theory and evidence from the UK

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Abstract

This paper presents a testable theoretical framework that extends the standard demand-side approach to modeling government expenditure on goods and services. The focus is on the adjustment of expenditure to disequilibria: we investigate whether the adjustment of UK exhaustive government expenditure between 1966 and 2002 to its long-run equilibrium path is symmetric. The evidence points to asymmetric adjustment to the demands of a representative voter over the election cycle but not between Labour and Conservative governments. Convergence to equilibrium is found to be faster during the later stages of each election cycle.

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1. Introduction

After its election victory in 1997 Britain's New Labour Government embarked on major reforms to the administration of fiscal policy. The government legislated for a set of principles to be applied when conducting fiscal policy. The principles are the backbone of the Code for Fiscal Stability[1]. One of the main arguments for reforms of this type is the need for macroeconomic policy to be sheltered from the myopic motives of political agents. The purpose of the present paper, in this context, is to investigate how political factors have shaped the long-run relationships describing the UK government's provision of goods and services.

We use the standard demand-side approach to modelling government expenditure that Borcherding (1985) described as the 'fiscal everyman approach'. This approach, following the Downsian tradition, is based on targeting the representative voter-taxpayer's preferences for publicly provided goods and services and has been used in previous empirical investigations of UK government expenditures by Tridimas (1992) and Ashworth (1995). These latter studies show that the preferences of the representative voter-taxpayer have had an impact on government expenditure growth but cannot tell us whether governments' delivery of goods and services to the representative voter varies across the election cycle. We consider how election cycles affect the way in which incumbent governments target the representative voter-taxpayer's preferences. Our concern is with the impact on UK government expenditures of the preferences of the representative voter-taxpayer. We consider whether these preferences are consistently targeted across the electoral cycle and irrespective of the incumbent's partisan (or ideological) persuasion. This

[1] For more details and an articulation of the Government's perspective on its policy-making reforms attention is drawn to HM Treasury (2002).

departs from the reaction-function approach to modelling the impact of elections on expenditure behaviour[2].

We outline a theoretical framework that focuses on the adjustment of expenditure towards its long-run equilibrium, that is, the representative voter-taxpayer's preferences. We consider the possibility that an asymmetry in the adjustment to equilibrium could result from either the time elapsed in the election cycle or the political parties. The empirical analysis is conducted using the cointegration tests with threshold adjustment as advanced by Enders and Siklos (2001). We use quarterly data to investigate adjustment to the equilibrium over the period from 1966 to 2002. The analysis focuses on exhaustive government expenditure[3] when considering the long-run equilibrium relationship for government goods and services in the United Kingdom.

The present analysis extends previous research in three respects. Firstly, we extend the standard demand-side approach to account for the effect of the time elapsed in the election cycle and political ideology on expenditure policy. We are able to assess whether the representative voter-taxpayer's preference is paramount throughout the election cycle. Secondly, we focus on the adjustment of expenditure to disequilibria. In particular, if the adjustment to the long-run equilibrium is asymmetric. Expenditure policy, in a dynamic context, is modelled as a decaying deviation from the representative voter-taxpayer's preferences. We consider the nature of any persisting deviation and its causes. Is it the result of opportunistic or partisan behaviour? Finally, it focuses exclusively on the provision of goods and services by the UK government. Section 2 of the paper outlines the theoretical

[2] This approach assesses the impact of elections on the responsiveness of government expenditures to a series of macroeconomic variables (see Tellier, 2003, for a recent survey). In earlier research (Easaw and Garratt, 1999, 2000), we adopted this approach when investigating UK expenditure cycles.
[3] Brown and Jackson (1990) use the adjective 'exhaustive' to differentiate between those expenditures that represent a claim on society's resources (purchase of inputs) from transfers, which represent a redistribution of resources.

framework. This forms the basis of the empirical investigation in Section 3. It includes the results from the residual-based approach to cointegration under asymmetric adjustment presented by Enders and Siklos (2001) as well as estimates of asymmetric Error Correction Models (ECMs). Finally, the main conclusions are drawn in Section 4.

2. Theoretical framework

The standard demand-function approach to modeling government expenditure on goods and services was initially designed to investigate the determination of local goods and services, but later adapted to national government expenditures.[4] The demand-function approach is based on a representative voter-taxpayer's set of preferences for government goods and services.[5] Niskanen (1978) argues that the preference set can be depicted as

$$lnQ = A + \eta lnR + \alpha lnY \qquad (1)$$

where Q is the quantity of government goods and services consumed by the average voter, R the perceived price paid and Y their real income[6].

Since Q, the amount captured by the representative voter-taxpayer, is unobservable it is necessary to use information about the quantity of goods and services produced for the population as a whole. The representative voter-taxpayer's consumption of government goods and services is then determined as follows

$$Q = \frac{X}{N^{\varphi}} \qquad (2)$$

[4] See Brown and Jackson (1990, Chapter 5) for a summary of the findings from early empirical estimations of the demand for local services.
[5] For a presentation of the representative-voter approach to the modelling of social benefits, see Boadway and Wildasin (1989).
[6] Niskanen also included an unspecified variable to capture autonomous conditions affecting the demand for government goods and services. We do not include this variable here as any proxy would be ad hoc.

where X is the amount of the good or service provided for the population, N the size of the population and ϕ measures the rivalry in consumption of government goods and services, sometimes referred to as their ‘publicness’.

The tax price per unit of government provision consumed by the representative voter when there are N voter-taxpayers can be written as

$$R = \frac{CX}{NQ} * \frac{T}{E} \tag{3}$$

where C is the average cost of the X units of the goods and services provided by government and T/E is the tax ratio measuring the proportion of this expenditure paid for out of taxation. Buchanan and Wagner (1977) argue that budget deficits increase government spending because they reduce the perceived price of government goods and services to the current generation. This is justified in several ways, for example by assuming that voters heavily discount future tax liabilities or lack awareness of the extent of future tax liabilities. The latter issue was identified by Rogoff and Sibert (1988) in the context of the election cycle. Their model shows how governments in the run-up to national elections might attempt to signal their administrative competency by offering a bundle of goods and services that is not wholly affordable out of general taxation. Rational voters would not ordinarily support this, but with an appropriate lag structure the need for additional revenues only becomes apparent after votes have been cast.

Following equations (1) to (3), we can write the total quantity of government goods and services as

$$lnX = A + \alpha lnY + \eta lnC + \eta ln(T/E) + ((\varphi - 1)\eta + \varphi)lnN \tag{4}$$

Equation (4) is now the aggregate demand for government goods and services. When using national accounts data for empirical estimation the equivalent to X is exhaustive government expenditure (*EX*) deflated by its cost (*C*). As Tridimas (1992) notes empirical work is interested in the real demand for government goods and services. Therefore, the income variable in equation (4) is the representative voter's real gross income, *y*, while the real cost of the government's provision of goods and services is the cost index of government provision relative to the index of prices, *C/P*. Finally, in abstracting in the presence of the provision of transfers, the taxation-financing variable is defined as taxation receipts net of transfers (NT) relative to exhaustive expenditure (EX). Therefore, the aggregate demand for government goods and services can be written as

$$ln(EX/C) = A + \alpha lny + \eta ln(C/P) + \eta ln(NT/EX) + ((\varphi - 1)\eta + \varphi)lnN \qquad (5)$$

Consequently, abstracting from the implied constraints on the parameters, we can estimate equation (6) and, hence, derive the long-run equilibrium of exhaustive expenditure. Wald Tests can then be conducted to test the restrictions on the parameters.

$$ln(EX/C) = \beta_0 + \beta_1 lny + \beta_2 ln(C/P) + \beta_3 ln(NT/EX) + \beta_4 lnN \qquad (6)$$

We now consider how short-run dynamics may affect this long-run equilibrium. To ease notation we use Z_t to depict the representative voter-taxpayer's preferences and, hence, Equation (6) is re-specified as follows

$$ln(EX/C)_t = \beta_0 + \beta lnZ_t \qquad (7)$$

Any divergence between government goods and services and the representative voter-taxpayer's preferences is assumed to be mean-reverting

$$ln(EX/C)_t - \beta_0 - \beta lnZ_t = \lambda[ln(EX/C)_{t-1} - \beta_0 - \beta lnZ_{t-1}] + \tau \qquad |\lambda| < 1 \qquad (8)$$

where $\tau/(1-\lambda)$ denotes the mean. From equations (7) and (8) we are able to derive the dynamics of the relationship between government goods and services and the representative voter-taxpayer's preferences (see Appendix 1 for the derivation)

$$\Delta(ln(EX/C)_t - \beta lnZ_t) = \gamma[ln(EX/C)_{t-1} - \beta_0 - \beta lnZ_{t-1}] \tag{9}$$

where $\gamma = (\lambda - 1)$. This suggests they co-move or are cointegrated over the long-run.

The adjustment (γ) may be asymmetric. The asymmetry could arise due to either partisan (or ideological) reasons or proximity to elections[7]. Incumbents of different political persuasion may vary in the way they choose to adhere to the representative voter-taxpayer when providing goods and services. This could be driven by ideological motivations. Likewise, incumbents, regardless of their ideological persuasion, may choose to conform to the representative voter-taxpayer differently depending on the time elapsed in the election cycle. Opportunism may lead governments to adhere more closely to the representative voter's preferences in the run up to elections (the pre-election period) than in the period following elections (the post-election period). Hence, the speed of adjustment of government expenditures to equilibrium in the pre-election period may be faster that in the pre-election period.[8] The test applied to opportunistic theory focuses on the impact of the election cycle.

We account for any asymmetric mean-reverting behavior, using a general specification as follows

[7] See Rogoff and Sibert (1988) and Rogoff (1990) for specific opportunistic models, Hibbs (1992) for a review of the early developments in partisan theory and Andrikopoulos, Loizides and Prodromidis (2004) and references therein for a recent review of the political business cycle literature as well as empirical evidence relating to fiscal policy in the nations of the EU.

[8] Because UK governments can call an election at any time, provided it is within 5 years of taking office, the empirical analysis takes the post-election period to be of pre-determined length. Hence, the pre-election period will vary in length.

$$\Delta[ln(EX/C)_t - \beta lnZ_t] = \gamma_1 I_t [ln(EX/C)_{t-1} - \beta_0 - \beta lnZ_{t-1}] + \gamma_2(1 - I_t)[ln(EX/C)_{t-1} - \beta_0 - \beta lnZ_{t-1}] \quad (10)$$

where $\gamma_1 \neq \gamma_2$ and I_t is the Heaviside indicator function. Adjustment to the long-run equilibrium may be either state-varying or depend on directional issues. In the case of state-varying behavior, the asymmetric relationship can be specified as a threshold autoregressive (TAR):

$$I_t = \begin{cases} 1 \text{ if } [ln(EX/C)_{t-1} - \beta_0 - \beta lnZ_{t-1}] \geq 0 \\ 0 \text{ if } [ln(EX/C)_{t-1} - \beta_0 - \beta lnZ_{t-1}] < 0 \end{cases} \quad (11)$$

The adjustment depends on whether government expenditure is above or below the level that depicts the representative voter-taxpayer's preferences. One may expect faster adjustment when government expenditure is below the representative voter-taxpayer's preferences. Governments are more likely to rectify any shortfalls when meeting the representative voter-taxpayer's preferences.

The directional issues are accounted for using a momentum threshold autoregressive (M-TAR)

$$I_t = \begin{cases} 1 \text{ if } [ln(EX/C)_{t-1} - \beta_0 - \beta lnZ_{t-1}] \geq [ln(EX/C)_{t-2} - \beta_0 - \beta lnZ_{t-2}] \\ 0 \text{ if } [ln(EX/C)_{t-1} - \beta_0 - \beta lnZ_{t-1}] < [ln(EX/C)_{t-2} - \beta_0 - \beta lnZ_{t-2}] \end{cases} \quad (12)$$

If in the last period government spending rose relative to its equilibrium then the change in any divergence is positive. The change in the divergence is negative if there was a fall in government spending relative to its equilibrium in the last period. We assume the speed of adjustment differs between these two cases. One may expect greater momentum when government expenditure has risen. Hence, expenditure is downwardly rigid. The remainder of the paper investigates empirically the implications of the theoretical model. In particular it is concerned with whether the adjustment to the long-run equilibrium is affected by elections or political persuasion and, hence, opportunism or partisan behavior.

3. Empirical analysis

Exhaustive government expenditure is expenditure on goods and services. It is the sum of the government's final consumption expenditure and its capital expenditure. Between 1966Q1 and 2002Q1 the average growth in real government expenditure on goods and services was 1.5% per annum.[9] Figure 1 shows the level and annual growth rate of real exhaustive government expenditure over this period[10]

Figure 1 [about here]

Figure 2 shows the ratio of the exhaustive government expenditure cost index[11] to the GDP deflator alongside the percentage of exhaustive expenditure that can be financed out of net taxation. The relative cost of government expenditure is seen to have increased over the period while the budgetary position has been more variable, but with the mid-1990s particularly notable for the weakness of the public finances when net taxation receipts covered only around 70% of total exhaustive expenditures.

Figure 2 [about here]

3.1 Testing for Cointegration with TAR and MTAR Adjustment

The augmented Dickey-Fuller (ADF) test is first used to determine the order of integration of the variables in equation (6). The results are summarized in Table 1. All variables are found to be I(1), which allows us to proceed with a residual-based approach to testing for cointegration.

[9] The average growth in total managed expenditure, which includes transfer payments, was 2.3% over the same period.
[10] All data is taken from Statbase provided by National Statistics and is available on-line at http://www.nationalstatistics.gov.uk/statbase/tsdintro.asp

Table 1 [about here]

The first step involves the estimation by LS of the long-run relationship for exhaustive government expenditure. Over the sample period 1966Q1-2002Q1 the following long-run equilibrium relationship is obtained (t-statistics in parentheses)

$$\underset{}{ln(EX/C)_t} = \underset{(4.002)}{5.5499} + \underset{(10.200)}{0.39383}\, ln y_t + \underset{(8.460)}{0.46638}\, ln(C/P)_t - \underset{(-10.564)}{0.19128}\, ln(NT/EX)_t + \underset{(1.151)}{0.47448}\, ln N_t \qquad (13)$$

All the expenditure elasticities are found to be statistically significant with the exception of the population elasticity.[12] The tax financing elasticity is seen to be significantly negative so that deficit financing increases the provision of good and services. Contrary to theory, but consistent with the finding of Tridimas (1992), increases in the cost of provision result in increased provision. Consequently, we are able to reject to the null hypothesis relating to equation (6) that $\beta_2 = \beta_3$.[13] But, further, we also reject the null hypothesis that the sum of the tax financing and cost elasticities is zero.[14] Increases in the effective tax price have not deterred provision perhaps because the complexity of the taxation system makes it difficult for the public to accurately infer the effective tax price or because there are complementarities in consumption between goods and services provided privately and by government. Using the residuals we are able to estimate a model of the form

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[11] The cost index of exhaustive government expenditure is the weighted average of the deflators for government final consumption and investment.

[12] As can be seen below the estimated long-run relationship in the absence of population is fundamentally unchanged. The empirical analysis presented later in this paper was also conducted in the absence of population again with no substantive differences found. These results are available on request from the authors.

$$ln(EX/C)_t = \underset{(4.002)}{7.1399} + \underset{(28.808)}{0.43474}\, ln y_t + \underset{(8.381)}{0.45462}\, ln(C/P)_t - \underset{(-11.394)}{0.19742}\, ln(NT/EX)_t$$

[13] The Wald Statistic for $\beta_2 = \beta_3$ is 173.7248 (p=0.000)

[14] The Wald Statistic on the restriction that the sum of the tax financing and cost elasticities is zero, i.e. $\beta_2 = -\beta_3$, is 17.8303 (p=0.000).

$$\Delta u_t = \beta_1 u_{t-1} + \sum_{i=1}^{P} \gamma \Delta u_{t-i} \tag{14}$$

Table 2 below gives the results of the residual-based cointegration test using both the Engel-Granger and TAR models.[15]

Table 2[about here]

The Akaike Information Criterion (AIC) suggests that a model using a single lagged change is appropriate, as shown in Table 2. The t-statistic for the coefficient on u_{t-1} is -3.7222. We are close to rejecting the null hypothesis of no cointegration at the 10% level.[16] But, this co-integration test is misspecified if adjustment to equilibrium is asymmetric. Following the method of Enders and Siklos (2001) we test for the possibility of asymmetric adjustment.

Again using the residuals we estimate a Threshold Autoregressive (TAR) model which takes the following form

$$\Delta u_t = \beta_1 I_t u_{t-1} + \beta_2 (1 - I_t) u_{t-1} + \sum_{i=1}^{P} \gamma \Delta u_{t-i} \tag{15}$$

where I_t is the Heaviside indicator function such that

$$I_t = 1\text{, if } u_{t-1} \geq 0$$

$$I_t = 0\text{, if } u_{t-1} < 0 \tag{16}$$

When the threshold value is taken to be zero, $\beta_1 I_t u_{t-1}$ is the adjustment if last period's residual is positive meaning that the government's provision of goods and services is above the equilibrium level. The adjustment is $\beta_2 (1 - I_t) u_{t-1}$ if last period's residual is negative and the government's provision is below the equilibrium level. From the third column of Table 2 we see that the point estimates of β_1 and β_2 indicate convergence. This suggests that we

[15] The presentation of results follows that used by Enders and Siklos (2001).
[16] The critical Engle-Granger value at the 10% significance level is –3.83.

can use the F statistic, the test value of which is denoted as F_C. The critical values reported in Enders and Siklos (2001) for the inclusion of a single lagged change indicate that we can reject the null hypothesis that $\beta_1 = \beta_2 = 0$ at the 1% significance level (the 1% critical value is approximately 8.30). Since a cointegrating relationship exists the null hypothesis of symmetric adjustment $\beta_1 = \beta_2$ can be tested using the Wald Test. The sample F value, denoted as F_A, is 0.14641 and has a *p* value of 0.701, which means that at conventional significance levels we are unable to reject the null hypothesis of symmetric adjustment.

We now consider whether the threshold depends on the previous period's change in the residual. The implication is that the residuals have greater momentum in a particular direction. Hence, such models are referred to as momentum-threshold autoregressive (M-TAR) models. The Heaviside indicator is set according to

$$I_t = 1, \text{ if } \Delta u_{t-1} \geq 0$$

$$I_t = 0, \text{ if } \Delta u_{t-1} < 0 \tag{17}$$

If in the last period government spending rose relative to its equilibrium then the change in the residual is positive. In this case adjustment in the latest period is $\beta_1 I_t u_{t-1}$. If, however, there was a fall in government spending relative to its equilibrium in the last period the change in the residual is negative and the adjustment in the latest period will be $\beta_2(1-I_t)u_{t-1}$. The fourth column of Table 2 reports the results of the M-TAR estimation. The point estimates of β_1 and β_2 again indicate convergence. The F_C value of 16.2296 allows us to reject the null hypothesis of no cointegration at the 1% level. Although the point estimates imply a greater persistence of disequilibria where government spending falls below its long-run equilibrium, the F_A of 1.3302 means that we fail to reject the null hypothesis of symmetric adjustment at the 10% level.

Following the method of Chan (1993) we are able to find the consistent estimate of the threshold parameter for the MTAR model. This is the threshold value that results in the lowest residual sum of squares. It is found to be -0.00265.[17] The estimated M-TAR model using this consistent estimate is reported in the last column of Table 2. The F_C value of 17.4057 indicates that we are able to reject the null hypothesis of no cointegration. The point estimates suggest convergence such that the speed of adjustment for positive discrepancies, where $\Delta u_{t-1} \geq 0.00265$, is more rapid than for negative discrepancies, where $\Delta u_{t-1} < 0.00265$. The larger of the two t-statistics is -0.604, which means that we cannot reject the null of no cointegration using what Enders and Siklos (2001) refer to as the t-Max(M) test. But, in the paper they show that the t-Max(M) test has 'low power' and so conclude that they 'cannot recommend using the t-Max(M) test' (p.172).[18] Therefore, since we are able to reject the null hypothesis of no cointergation we can test the null hypothesis of symmetric adjustment using the Wald Test.[19] The F_A value of 3.2692 has a p value of 0.071, which means that we are able to reject the null hypothesis of symmetric adjustment at the 10% level.

[17] This equates to close on £1 billion (2001 prices) per quarter. During 2001, a change of this magnitude was equivalent to 2.1% of expenditure.

[18] Enders and Siklos (2001) note that the situation is quite different in the TAR model where 'the power of the t-Max statistic is usually superior' to the FC statistic 'when the size of the test is 1% and there is a reasonable amount of asymmetry' (p.171).

[19] Enders and Siklos (2001) face the same situation when analysing the Federal Funds Rate.

Following Enders and Siklos (2001, p. 174) who based their model selection on the AIC, we find that the momentum-consistent TAR model fits the data better than the other models in Table 2. Hence, there is greater momentum for quarter-on-quarter increases in exhaustive expenditures of more than £1 billion (2001 prices), relative to the equilibrium level. It displays downward rigidity. However, some caution should be attached to this interpretation since it ignores both the ideological persuasion of the governing party and time elapsed in the election cycle. These are issues we now address.

Threshold Autoregressive Models allow us to consider whether the nature of adjustment to long-run relationships is state contingent. As outlined in the preceding section this approach is ideally suited to an analysis of the political system on government expenditure determination. We consider two possible ways in which politics might impact on fiscal policy. The first is a partisan or ideological effect. The second, which has tended to dominate the literature, focuses on electoral opportunism. The application of partisan theory in the current context involves considering the null hypothesis that the speed of adjustment of UK government expenditures during Labour (left of centre) and Conservative (right of centre) governments is the same. Consequently, the Heaviside indicator function is defined as

$I_t = 1$, if u_{t-1} occurred under a *Labour government*

$I_t = 0$, if u_{t-1} occurred under a *Conservative government* (18)

Table 3 [about here]

Table 3 below outlines the relevant results. The results from the TAR model, which accounts for the partisan effects are presented in the second column of the Table. The point estimates of β_1 (Labour) and β_2 (Conservative) are consistent with convergence and the F_C of 15.4162 indicates that the null hypothesis that $\beta_1 = \beta_2 = 0$ can be rejected at the 1%

level. However, the F_A value of 0.00051 means that we are unable to reject the null hypothesis of symmetric adjustment at conventional significance levels. The speed of adjustment of government spending on goods and services to equilibrium is found to have been the same under Labour and Conservative governments.

The test applied to opportunistic theory focuses on the impact of the election cycle. Specifically, it involves consideration of the null hypothesis that the speed of adjustment of government expenditures to equilibrium in a 'post-election period' of pre-determined length is the same as in the remainder of the election period or the 'pre-election period'. This leads us to re-define the Heaviside indicator function as

$I_t = 1$, if u_{t-1} occurred in a *post-election quarter*

$I_t = 0$, if u_{t-1} occurred in a *pre-election quarter* (19)

The variable election cycle length in the UK does present problems for empirical analysis of this sort. Hence, our results should be viewed as illustrative of possible differences in post and pre-election expenditure behaviour by the UK government.[20] We considered alternative lengths for the post-election period ranging from 4 to 8 quarters after each General (national) Election in the United Kingdom. The results are presented in Table 3. Of the alternative election cycle models, the AIC infers that the 6 quarter model type best fits the data. The point estimates of β_1 (post-election) and β_2 (pre-election) indicate convergence, while the F_C values allow us to reject the null hypothesis that $\beta_1 = \beta_2 = 0$. Moreover, the F_A value allows us to reject the null hypothesis of symmetric adjustment at the 1% level.

[20] The case of the two UK national elections held in 1974 (February and October) is one example of the problematic nature of applying the concepts of pre- and post-election periods to the UK. We have treated these 2 elections as separate events. Hence, the post-election periods of these 2 elections overlap because of their close proximity in time.

The evidence points to the adjustment to disequilibrium being slower in the early part of the UK election cycle. The turning point is around 18 months after the national election.

3.2 Asymmetric error correction models

Tables 4 and 5 present the estimates of the error correction models which enable us to see if our findings are corroborated after information regarding the short-run dynamics is accounted for.

Tables 4 and 5 [about here]

For completeness the estimates of the symmetric ECM are also reported in Table 4. The lag structure of the dynamic terms is determined by reference to the AIC. Across the models the short-run dynamics are statistically significant at the 10% level or lower with the exception of population whose long-run elasticity was also statistically insignificant. The short-run tax-financing elasticity is negative consistent with government using deficit financing to fund increased expenditure. In absolute terms the short-run tax-financing elasticity is larger than its long-run counterpart. The short-run cost elasticity is found be positive, as was the case with the long-run elasticity, and again contrary to expectation. But, the absolute value of the short-run tax-financing elasticity is found to be significantly greater than the short-run cost elasticity at the 10% level or lower. Therefore, in the short-run decreases in the effective tax price are attributable to increases in government spending on goods and services.

The error correction models in Table 4 largely corroborate the earlier findings such that we are again unable to reject the null hypothesis of symmetric adjustment at conventional significance levels for the TAR and MTAR models with zero thresholds. But, the error correction terms in the MTAR model using the consistent threshold estimate are found to be significantly different at the 10% level. Convergence to equilibrium is faster for positive discrepancies from the threshold. Decreases in expenditures in excess of the consistent threshold (£1 billion, 2001 prices) are found to display greater persistence, as detected from the analysis of the residuals of the equilibrium expenditure relationship. Furthermore, based on the AIC, the MTAR error-correction model with the consistent estimate of the threshold is found to better fit the data than the TAR and M-TAR models. This too is consistent with those results from an analysis of the residuals.

Table 5 confirms that there is no difference between Labour and Conservative governments in the speed with which exhaustive government expenditure converges on the long-run equilibrium relationship. However, the most significant finding is that the asymmetric adjustment of government expenditures to equilibrium across the election cycle is also evident once the appropriate error correction model is estimated. The AIC infers that the preferred political business cycle model, having incorporated the dynamics, has a turning point around 15 months after national elections, just a little earlier than was suggested by the analysis of the residuals. Up to this point disequilibria in the level of provision of goods and services have a significantly greater persistence than later in the election cycle. As the election cycle unfolds and the next election draws closer exhaustive government expenditure converges more quickly on equilibrium and so the representative voter-taxpayer's preferences. The inference is that governments of all persuasions attribute increasing importance to preferences of the representative voter as an election approaches.

Using the AIC[21] to compare across the ECM models in Tables 4 and 5 we find that the preferred model of exhaustive government expenditure in the United Kingdom is the

[21] The choice of the preferred ECM based on the AIC follows the approach of Enders and Siklos (2001, p.175).

opportunistic political business cycle model. Hence, the preferred model incorporates an asymmetric adjustment to disequilibria that is contingent on the point in the election cycle.

3.3 Discussion of results

The results provide further evidence that the long-run equilibrium of government expenditure in the United Kingdom is shaped by the preferences of the representative voter. Evidence of a co-integrating relationship between exhaustive expenditures and a series of demand variables demonstrates the importance of the representative voter on policy-making. Therefore, the long-run equilibrium can be labelled as Downsian since Anthony Downs (1957) predicted that political competition and the need to maximise votes would lead political parties to identify the preferences of the representative voter.

Using threshold cointegration models we observe that the adjustment of exhaustive expenditures in the UK to the preferences of the representative voter exhibits asymmetric properties. The results show that this asymmetry occurs when converging on its long-run equilibrium; governments do not lose sight of their need to appeal to the representative voter at general elections.

While the momentum of divergence can be shown to matter when modelling exhaustive government expenditure a better model fit is obtained when the residuals are assigned to one of two points on the election cycle. The adjustment of UK exhaustive government expenditure is significantly slower in a post-election period of between 15 and 18 months after each general election cycle. The comparative speeds of adjustment in the post and pre-election periods can be viewed as indicators of the importance the government attaches to the preferences of the representative voter in these periods. The faster the speed of adjustment the more important is the representative voter. While expenditure convergence is observed across the election cycle the representative voter's preferences matter most in the latter part of the UK general election cycle, that is, prior to the next election.

UK governments afford themselves most discretion in the earlier stages of an election cycle. Concerns about re-election are at their most distant and so convergence to the representative voter's preferences is slower. However, there is no evidence to suggest that convergence is contingent on the political persuasion of the ruling government. Clearly both Labour and Conservatives are equally concerned about meeting the representative voter's preferences. The discretion observed is opportunistic, borne out of the election cycle rather than political orientation.

4. Concluding remarks

This paper supports the political-economy view that account should be taken of political variables when modeling government expenditures. Our contribution has been to offer new insights into how politics affects public spending in the United Kingdom. We have set out a testable theory of expenditure policy as a decaying deviation from the representative voter-taxpayer's preference, which may be asymmetric. Using cointegration analysis with threshold adjustment, we have assessed the nature and cause of the divergence. We have found evidence that adjustment to equilibrium differs according to the direction of momentum and is contingent on the time elapsed in the election cycle.

The detection of an asymmetric exhaustive expenditure policy has important implications for the way in which we model government expenditures. Cointegration tests are misspecified if adjustment is asymmetric and asymmetries necessitate the estimation of asymmetric error correction models. In assessing asymmetric expenditure policy, we find a more rapid adjustment to long-run equilibrium for positive changes where there has occurred a quarter-on-quarter increase in exhaustive expenditures of more than £1 billion (2001 prices) relative to the equilibrium level. We have also been able to attribute policy asymmetry to the point in the election cycle. The results indicate that there is greater persistence in the post-election period, while convergence of expenditure on equilibrium levels occurs more quickly in the run-up to elections. Model selection criteria favour the estimation of an asymmetric error correction model for exhaustive government expenditure in the United Kingdom, where the asymmetry arises out of the election cycle with a turning point around 15 months after the national election.

Our conclusions support the idea that the representative voter-taxpayer plays an important role in influencing the UK government's conduct of its exhaustive expenditure policy. Policy decisions are made that enable expenditure levels to converge on equilibrium. However, the uncovering of asymmetric adjustment suggests that UK government affords itself some discretion in the provision of goods and services.

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Appendix A1

The long-run relationship is:

$$\ln(EX/C)_t = \beta_0 + \beta \ln Z_t \qquad [A.1]$$

with mean-reverting divergence:

$$\ln(EX/C)_t - \beta_0 - \beta \ln Z_t = \lambda[\ln(EX/C)_{t-1} - \beta_0 - \beta \ln Z_{t-1}] \qquad |\lambda| < 1 \qquad [A.2]$$

where the mean is zero. Re-arranging the terms

$$\ln(EX/C)_t = \beta_0 + \beta \ln Z_t - \lambda\beta_0 - \lambda\beta \ln Z_{t-1} + \lambda \ln(EX/C)_{t-1} \qquad [A.3]$$

Adding and subtracting βlnZ_{t-1}:

$$\ln(EX/C)_t = (1-\lambda)\beta_0 + \beta\Delta \ln Z_t + (1-\lambda)\beta \ln Z_{t-1} + \lambda \ln(EX/C)_{t-1} \qquad [A.4]$$

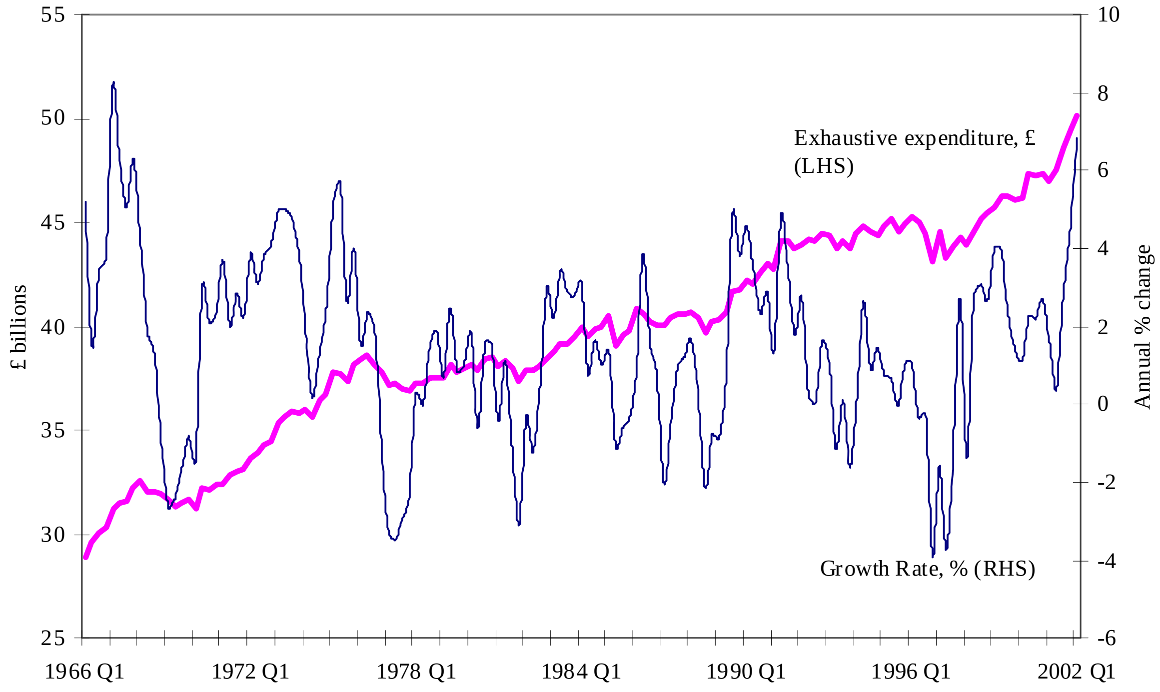
Subtracting $\ln(EX/C)_{t-1}$ from both sides;

$$\Delta \ln(EX/C)_t = (1-\lambda)\beta_0 + \beta\Delta \ln Z_t + (1-\lambda)\beta \ln Z_{t-1} - (1-\lambda) \ln(EX/C)_{t-1} \qquad [A.5]$$

or alternatively,

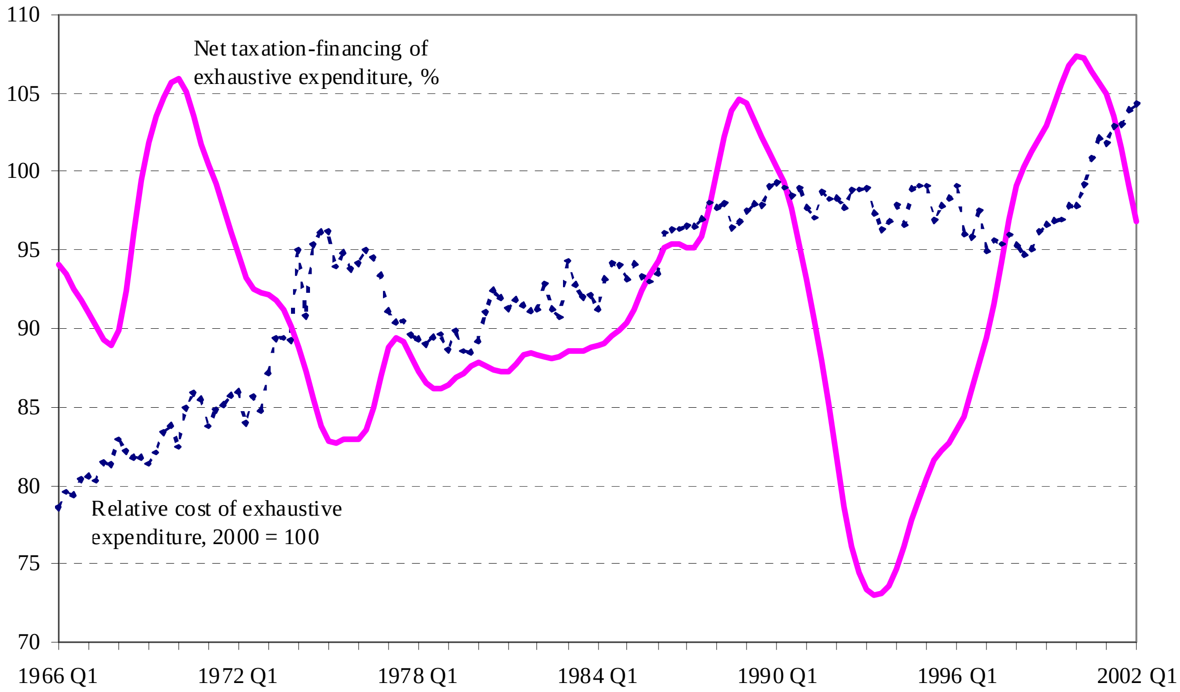
$$\Delta \ln(EX/C)_t = (\lambda - 1)[\ln(EX/C)_{t-1} - \beta_0 - \beta \ln Z_{t-1}] + \beta\Delta \ln Z_t \qquad [A.6]$$

Figure 1: UK exhaustive government expenditure, level and annual growth rate, 1966Q1 – 2002Q1, quarterly



Source: Blue Book, Office for National Statistics

Figure 2: Percentage of exhaustive expenditure financed out of net taxation and relative cost of exhaustive expenditure, 1966Q1 – 2002Q1, quarterly



Source: Blue Book, Office for National Statistics

Table 1: Augmented Dickey-Fuller tests
(sample 1966Q1 – 2002Q1, quarterly)

Variable	**ADF(*p,t*) I(1)/I(0)**	**ADF(*p,t*) I(2)/(1)**
ln(EX/C)	-2.6456 (1,1)	-13.8674 (0,0)
ln(C/P)	-1.7410 (1,0)	-15.7397 (0,0)
Lny	-2.1836 (1,1)	-12.3249 (2,0)
ln(NT/EX)	-2.4388 (6,0)	-4.5747 (5,0)
lnN	0.00452 (3,0)	-4.5464 (2,0)

ADF test includes a constant. Null hypothesis is non-stationary series. p is the number of lagged changes used in the ADF test and t indicates whether a time trend is included. SBC is used in determining preferred specification. The 5% critical value in the absence of a time trend is -2.8815 and -3.4415 in the presence of a time trend.

Table 2: Estimates of exhaustive government expenditure residuals (sample 1966Q1 – 2002Q1, quarterly)

	Symmetric convergence	Asymmetric convergence		
	Engle-Granger	Threshold	Momentum	Momentum-consistent
β_1 [a]	-0.20609 (-3.722)	-0.18130 (-2.124)	-0.26803 (-3.481)	-0.26890 (-4.1372)
β_2 [a]	NA	-0.22229 (-3.184)	-0.14689 (-1.949)	-0.059199 (-0.604)
γ_1 [a]	-0.21991 (-2.746)	-0.22172 (-2.756)	-0.22449 (-2.803)	-0.20440 (-2.558)
AIC	440.1683	439.2430	439.8476	440.8186
F_C [b]	NA	15.5053	16.2296	17.4057
F_A [c]	NA	0.14641 (0.702)	1.3301 (0.249)	3.2692 (0.071)

[a] In parentheses are the t-statistics
[b] F-statistic for null hypothesis that the coefficients on the regressors are zero (no cointegrating relationship)
[c] Wald Statistic for null hypothesis that $\beta_1 = \beta_2$ (symmetric adjustment). *P* value in parentheses

Table 3: Estimates of exhaustive government expenditure residuals (sample 1966Q1–2002Q1, quarterly)

	Partisan	Election cycle				
		4 quarters	5 quarters	6 quarters	7 quarters	8 quarters
β_1 [a]	-0.20555	-0.064312	-0.045591	-0.061408	-0.11876	-0.14120
	(-2.194)	(-0.722)	(-0.558)	(-0.816)	(-1.662)	(-2.044)
β_2 [a]	-0.20637	-0.28728	-0.32359	-0.34855	-0.32021	-0.30780
	(-3.029)	(-4.226)	(-4.602)	(-4.666)	(-3.944)	(-3.603)
γ_1 [a]	-0.21981	-0.19624	-0.19045	-0.18606	-0.19313	-0.20045
	(-2.693)	(-2.450)	(-2.402)	(-2.348)	(-2.397)	(-2.485)
AIC	439.1684	441.2177	442.6087	442.9761	441.0021	440.3964
F_C [b]	15.4162	17.8937	19.6158	20.0761	17.6297	16.8924
F_A [c]	0.00051	4.0712	6.9008	7.6573	3.6374	2.4258
	(0.994)	(0.044)	(0.009)	(0.006)	(0.056)	(0.119)

[a] In parentheses are the t-statistics
[b] F-statistic for null hypothesis that the coefficients on the regressors are zero (no cointegrating relationship)
[c] Wald Statistic for null hypothesis that $\beta_1 = \beta_2$ (symmetric adjustment). *P* value in parentheses

Table 4: Estimates of error correction model (sample 1966Q1–2002Q1, quarterly)

	Symmetric	Asymmetric		
		Threshold	Momentum	Momentum-consistent
Constant	0.0020538 (1.688)*	0.0014703 (0.891)	0.0022999 (1.841)*	0.0022620 (1.864)*
$\Delta \ln(EX/C)_{t-1}$ [a]	-0.18110 (-2.321)**	-0.18231 (-2.330)**	-0.18541 (-2.370)**	-0.16328 (-2.091)**
$\Delta \ln y_t$ [a]	0.23261 (2.536)**	0.23299 (2.534)**	0.23422 (2.551)**	0.22912 (2.516)**
$\Delta \ln(C/P)_t$ [a]	0.13006 (1.962)**	0.12746 (1.913)*	0.13454 (2.022)**	0.14045 (2.126)**
$\Delta \ln(NT/EX)_t$ [a]	-0.29074 (-4.530)**	-0.29148 (-4.528)**	-0.28849 (-4.487)**	-0.29946 (-4.686)**
$\Delta \ln N_t$ [a]	1.2542 (1.075)	1.1844 (1.006)	1.1517 (0.981)	1.1730 (1.012)
u_{t-1} [a]	-0.18288 (-3.545)**			
$I_t u_{t-1}$ [a]		-0.13319 (-1.234)	-0.22895 (-3.120)**	-0.24152 (-3.949)**
$(1-I_t)u_{t-1}$ [a]		-0.22069 (-2.488)**	-0.13984 (-1.970)**	-0.04842 (-0.525)
AIC	448.9954	448.1409	448.4071	449.6038
F [b]	8.1309	6.9720	7.0697	7.5131
F_A [c]		0.27520 (0.377)	0.78020 (0.377)	3.0733 (0.080)

[a] In parentheses are the t-statistics (** significant at 5% level, * significant at 10% level)
[b] F-statistic for null hypothesis that the coefficients on the regressors are zero.
[c] Wald Statistic for null hypothesis that coefficient on $I_t U_{t-1}$ equals the coefficient on $(1-I_t)U_{t-1}$. *P* value in parentheses

Table 5: Estimates of error correction model (sample 1966Q1–2002Q1, quarterly)

	Partisan	Election cycle				
		4 quarters	5 quarters	6 quarters	7 quarters	8 quarters
Constant	0.0020761	0.0020128	0.0019752	0.0019104	-0.0019019	-0.0019265
	(1.694)	(1.681)	(1.667)*	(1.605)	(1.576)	(1.587)
$\Delta\ln(EX/C)_{t-1}$ [a]	-0.12800	-0.14770	-0.15272	-0.15066	-0.15325	-0.16215
	(-1.906)*	(-1.892)*	(-1.996)**	(-1.955)**	(-1.953)**	(-2.063)**
$\Delta\ln y_t$ [a]	0.23448	0.24506	0.25690	0.25783	0.24850	0.24165
	(2.537)**	(2.711)**	(2.866)**	(2.862)**	(2.727)**	(2.642)**
$\Delta\ln(C/P)_t$ [a]	0.12800	0.12419	0.12768	0.14020	0.13521	0.13726
	(1.906)*	(1.903)*	(1.980)**	(2.161)**	(2.060)**	(2.076)**
$\Delta\ln(NT/EX)_t$ [a]	-0.28929	-0.29326	-0.30074	-0.30504	-0.30817	-0.30879
	(-4.467)**	(-4.643)**	(-4.808)**	(-4.846)**	(-4.805)**	(-4.755)**
$\Delta\ln N_t$ [a]	1.2620	1.4177	1.4244	1.3796	1.2894	1.2814
	(1.077)	(1.233)	(1.253)	(1.209)	(1.116)	(1.103)
$I_t u_{t-1}$ [a]	-0.19855	-0.024895	-0.011268	-0.046614	-0.097010	-0.12213
	(-2.253)**	(-0.296)	(-0.147)	(-0.658)	(-1.447)	(-1.885)*
$(1-I_t)u_{t-1}$ [a]	-0.17448	-0.26892	-0.30168	-0.31375	-0.29418	-0.27893
	(-2.711)**	(-4.302)**	(-4.690)**	(-4.515)**	(-3.870)**	(-3.449)**
AIC	448.4071	450.8786	452.4743	451.8545	450.0336	449.2346
F [b]	7.0697	7.9935	8.6069	8.3671	7.6741	7.3755
F_A [c]	0.048285	5.5581	8.7304	7.4899	3.9061	2.3618
	(0.826)	(0.018)	(0.003)	(0.006)	(0.048)	(0.124)

[a] In parentheses are the t-statistics (** significant at 5% level, * significant at 10% level)
[b] F-statistic for null hypothesis that the coefficients on the regressors are zero
[c] Wald Statistic for null hypothesis that coefficient on $I_t U_{t-1}$ equals the coefficient on $(1-I_t)U_{t-1}$. *P* value in parentheses