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An Empirical Investigation into Government Spending and Private Sector Behaviour

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CONTENTS

ABSTRACT

We examine whether there is a significant relationship between government and private consumption for Canada. We derive estimating equations between the two types of consumption under both cointegration and no-cointegration assumptions. This distinction seems to have been largely ignored in previous work in the literature. Our results suggest that this distinction is an important one. For government spending on goods and services, our pretests do not allow us to firmly conclude whether government and private consumption are cointegrated or not. Therefore, we estimate the relationship under both cointegration and no cointegration assumptions. Under the cointegration assumption we find the two types of consumption to be complements, whereas under the no-cointegration assumption we find them to be substitutes. For government investment spending we are unable to find any evidence consistent with cointegration. Under the maintained assumption of no cointegration we find no statistically significant relationship but an economic relationship that implies they are complements rather than substitutes.

RÉSUMÉ

Dans la présente étude, les auteurs examinent s'il existe une relation significative entre les dépenses de consommation publiques et privées au Canada. À cette fin, ils ont dérivé les équations à estimer sous les hypothèses de cointégration et d'absence de cointégration entre les deux types de consommation. En général, cette distinction ne semble pas avoir été prise en compte dans les travaux effectués dans le domaine. Les résultats de l'étude donnent pourtant à penser que la distinction est importante. En ce qui a trait aux dépenses en biens et services du secteur public, les tests préliminaires menés par les auteurs ne permettent pas de conclure avec certitude ni à la cointégration ni à la non-cointégration des dépenses de consommation publiques et privées. C'est pourquoi les auteurs estiment la relation suivant chacune des deux hypothèses. Ils ont découvert, dans le premier cas, que les deux types de consommation sont complémentaires et, dans le deuxième, qu'ils sont substituables. Pour ce qui est des dépenses publiques d'investissement, les résultats obtenus n'ont pu étayer l'hypothèse de cointégration. Sous l'hypothèse d'absence de cointégration, ils n'ont constaté aucune relation statistique significative, mais une relation économique qui indique que les dépenses d'investissement et celles de consommation sont complémentaires plutôt que substituables.

1. INTRODUCTION

The hypothesis that government spending may substitute for private consumption is a relatively new one in economics. Bailey (1971) first proposed the idea that there may be a degree of substitutability between government and private consumption. Barro (1981) incorporated this hypothesis into a general model of consumption to allow for a direct effect of government purchases of goods and services on consumers' utility. Over the last decade there has been a growing literature examining the response of private consumption to changes in government spending. For example, Aiyagari, Christiano and Eichenbaum (1992) and Baxter and King (1993) explore the effect of government spending shocks on various economic aggregates in a one-sector neoclassical growth model with constant returns to scale and variable labour supply. These authors find, *inter alia,* that increases in government spending unambiguously lead to a decline in private consumption. Empirically, Kormendi (1983) and Aschauer (1985) work within the permanent-income framework and estimate a significant degree of substitutability between private and public consumption for the United States. Moreover, Ahmed (1986) examines the effects of U.K. government consumption in an intertemporal substitution model and finds that government expenditures tend to crowd out private consumption.

However, recent papers have called into question this negative comovement between government and private consumption. Devereux, Head and Lapham (1994) examine the impact of government spending shocks in a neoclassical model with increasing returns to scale and monopolistic competition. The key feature of their model is that an increase in government consumption generates an endogenous rise in aggregate productivity. The increase in productivity raises the real wage sufficiently that there is a substitution away from leisure and into consumption. Thus, an increase in government expenditures leads to an increase in private consumption. Empirically, Karras (1994) examines the response of private consumption to increases in government spending across a number of countries and finds that public and private consumption are better described as complementary rather than as substitutes. In sum, there appears to be no clear consensus among researchers on the response of private consumption to changes in government spending.

Given the wide range of empirical results, another investigation into this question seems warranted. This paper examines whether or not there is a relationship between government and private consumption for Canadian data. Our approach, however, differs from the previously mentioned empirical papers in one important way - we explicitly take into account the nonstationarity of the data. That is, we derive estimating equations between the two types of consumption under both cointegration and no-cointegration assumptions. Previous researchers have tended to ignore this important distinction.

The organization of this paper is as follows. Section 2 presents our model and derives equations to be estimated under both the cointegration and no-cointegration assumptions. Section 3 describes the data and presents results from our pretests. Section 4 reports the estimation results and attempts to address a concern recently noted by Graham (1993) regarding the robustness of the relationship between government and private consumption when real disposable income is included in the model. Section 5 concludes.

2. THE MODEL

Following Bailey (1971) we define effective consumption as

$$C_i^* = C_i + \alpha G_i \tag{1}$$

where C_i is real private expenditure, G_i is real government expenditure and α is the parameter measuring the relationship between C_i and G_i. Suppose that a representative consumer chooses consumption to maximize expected lifetime utility:

$$E_t\left[\sum_{i=t}^{T}\beta^{i-t}U(C_i^*)\right] \tag{2}$$

subject to the following period-by-period budget constraint:

$$A_{i+1} = (A_i + Y_i - C_i^* - (1-\alpha)G_i)(1+R) \tag{3}$$

where E_i is the expectations operator based on period i information, $\beta \in (0, 1)$ is a discount factor, A_i is real financial assets net real government debt at the beginning of period i, Y_i is real labour income in period i, and R is a time invariant real rate of interest. Finally, assume that U is increasing and concave in its arguments, and that $\partial U(0)/\partial C^* \to \infty$.

The Lagrangean for the problem is given by

$$E_0\left[\sum_{t=0}^{\infty}\beta^t\{U(C_t^*) - \Phi_t(A_{t+1} - (A_t - Y_t + C_t^* + (1-\alpha)G_t))(1+R)\}\right] \tag{4}$$

where Φ_t is the Lagrange multiplier associated with the budget constraint equation (3). The necessary first-order conditions for period t include the following equations

$$\partial U_t/\partial C_t^* = \Phi_t \tag{5}$$

$$E_t[\beta(1+R)\Phi_{t+1}] = \Phi_t \tag{6}$$

for $t = 1, 2, \ldots$, where $\partial U(C_t^*)/\partial C_t^* = \partial U_t/\partial C_t^*$.

The Euler equation between adjacent periods t and $t+1$ can be derived by substituting equation (5) for Φ_t and Φ_{t+1} into equation (6):

$$E_t[\beta(1+R_t)(\partial U_{t+1}/\partial U_t)] = 1 \tag{7}$$

To exploit the empirical implications of the model, we assume that the change in marginal utility is negligibly small from period to period so that equation (7) can be expressed as

$$E_t C^*_{t+1} = [\beta(1+R)]^{\sigma} C^*_t \tag{8}$$

where $\sigma = -U'(C^*)/C^* U''(C^*)$ is the intertemporal elasticity of substitution.

Equation (8) implies the following econometric relationship:

$$C^*_{t+1} = \gamma C^*_t + v_{t+1} \tag{9}$$

where $E_t[v_{t+1}] = 0$. When $\gamma = [\beta(1+R)]^{\sigma}$ is less than unity, it implies that C^*_t is a stationary or I(0) variable, that is, $C_t + \alpha G_t$ is an I(0) variable. This result suggests that if C_t and G_t are each I(1) or nonstationary, then they will be cointegrated in the sense of Engle and Granger (1987) with the cointegrating parameter given by α. On the other hand, when γ is equal to unity, and C_t and G_t are each I(1), then C_t and G_t will not be cointegrated.

Note that equation (9) can be expressed as

$$C_t + \alpha G_t = \gamma(C_{t-1} + \alpha G_{t-1}) + v_t \tag{10}$$

or more compactly as

$$C_t = \eta_1 C_{t-1} + \eta_2 G_t + \eta_3 G_{t-1} + v_t \tag{11}$$

where $\eta_1 = \gamma$, $\eta_2 = -\alpha$, $\eta_3 = \gamma\alpha$ and $v_t \sim NID(0, \sigma^2_v)$.

If $\gamma < 1$ such that C_t and G_t are cointegrated, least-squares (LS) estimation of equation (11) although T-consistent, will not be asymptotically optimal and will have non-standard limiting distributions that make statistical inference difficult to conduct. It is thus desirable to find a transformation of the model that yields asymptotically efficient estimates and allows us to perform inference on the estimated parameters using standard distributional theory. One such procedure has been proposed by Wickens and Breusch

(1988). The procedure yields an equation where the only variables in levels are those in the cointegrating regression, and all other variables are in first-difference form, *viz.*,

$$C_t = \lambda_1 G_t + \lambda_2 \Delta C_t + \lambda_3 \Delta G_t + \varepsilon_t \tag{12}$$

where $\lambda_1 = (\eta_2 + \eta_3) \mathbin{/} (1 - \eta_1)$, $\lambda_2 = -\eta_1 \mathbin{/} (1 - \eta_1)$, $\lambda_3 = -\eta_3 \mathbin{/} (1 - \eta_1)$ and $\varepsilon_t = v_t \mathbin{/} (1 - \eta_1)$ and is $NID(0, \sigma_\varepsilon^2)$. The resulting equation can then be estimated using instrumental variables (IV). Notice that the Wickens and Breusch procedure requires only one-step estimation. If the variables C_t and G_t are I(1), then the IV estimator of λ_1 from equation (12) is T-consistent. Moreover, this IV estimator has a mixed normal limiting distribution and the estimators of the η_i ($i = 1,2,3$) are asymptotically normal, which allows inference to proceed using standard asymptotic theory (see Amano and Wirjanto 1993).[1]

If $\gamma = 1$ such that C_t and G_t are not cointegrated, (11) reduces to a simple first-difference equation of the form

$$\Delta C_t = -\alpha \Delta G_t + v_t \tag{13}$$

where $E_{t-1}[v_t] = 0$ and which can be consistently estimated by IV.

3. DATA DESCRIPTION AND PRETEST RESULTS

From the perspective of the underlying permanent income model of effective consumption, we would prefer measures of government consumption that are separated along intertemporal lines. With this in mind we use two measures of real Canadian government consumption: (i) current government purchases of goods and services (GC)

1. There obviously are other efficient procedures which after some modification may be applied to obtain estimates of equation (11). Examples of these estimators include the three-step procedure of Engle and Yoo (1991), the full-information estimator proposed by Phillips (1991), the canonical cointegrating regression approach of Park (1992) and the modified LS estimators developed by Phillips and Hansen (1990), Phillips and Loretan (1991) and Stock and Watson (1993).

and (ii) total government investment (GI). Both measures are deflated by the total population of age 15 and over to give per capita variables. Similarly, per capita consumption series (C) is obtained by dividing real personal consumption of non-durable goods and services by the total population of age 15 and over. All series are seasonally adjusted, quarterly, span the sample period 1953Q1 to 1993Q2, and are used in logarithm form.[2] Further details of the data and their sources can be found in the Data Appendix.

We begin by examining the time-series properties of each of the series. To this end we use the augmented Dickey-Fuller (1979) test suggested by Said and Dickey (1984), the non-parametric Phillips and Perron (1988) Z_α test and a modification of the Z_α test proposed by Stock (1991) - the MZ_α test.[3] These tests allow us to formally test the null hypothesis that a series is I(1) against the alternative that it is I(0). The test statistics are reported in Table 1 (p. 15). For all three variables, the null hypothesis of a unit root cannot be rejected even at the 10 per cent level of significance. Therefore we conclude that the variables under consideration are well characterized as I(1) processes.

To examine whether there is evidence consistent with cointegration between private and government consumption, we begin with the two-step approach proposed by Granger (1983) and later refined by Engle and Granger (1987). Specifically, we employ the augmented Dickey-Fuller test suggested by Engle and Granger, the normalized bias version of the Phillips and Perron test proposed by Phillips and Ouliaris (1990) and MZ_α test developed by Stock (1991). We test the null hypothesis of no cointegration using two versions of the following test regression

$$c_t = a_0 + a_1 t + a_2 g_t + \vartheta_t \tag{14}$$

where lower case letters represent logs and g_t represents either government consumption

2. The period 1953 to 1993 is the longest span for which we were able to obtain data.
3. Stock's Monte Carlo evidence suggests that the MZ_α test has better finite sample properties in terms of both power and size relative to the other unit-root tests.

of goods and services or government investment expenditures. The first version is simply equation (19), while the second version sets a_1 equal to zero. Evidence of cointegration in the former suggests that the linear stationary combinations of the I(1) variables have a non-zero linear trend (stochastic cointegration). In contrast, evidence of cointegration in the latter corresponds to deterministic cointegration which implies that the same cointegration vector eliminates deterministic trends as well as stochastic trends.[4]

The results of the cointegration tests are reported in Table 2 (p. 15). For public consumption of goods and services, we see that the augmented Engle and Granger test admits evidence of deterministic cointegration at the 10 per cent level, while the MZ_α test statistic suggests the presence of stochastic cointegration at about the 12 per cent level. When g_t is set equal to government investment, we see that none of the tests are able to reject the null hypothesis of no cointegration. However, it is well known that in the presence of persistent roots, these tests will tend to lack power to detect a cointegrating relationship in the data, even when one is present. Thus, it is difficult to discern from the results whether the inability to reject the null hypothesis actually reflects a non-cointegrated system or simply the weak power of these cointegration tests. To control for this problem, we also apply recently developed tests that have cointegration as their null hypothesis.

Tests for the null of cointegration against the alternative of no cointegration have recently been proposed by Hansen (1992) and Shin (1994). The former is a series of parameter constancy tests for I(1) processes that can also be viewed as tests for the null of cointegration against the alternative of no cointegration. These are the *Lc*, *MeanF* and *SupF* tests. The Shin test is an extension of the Kwiatkowski, Phillips, Schmidt and Shin

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4. See Ogaki and Park (1989) for a discussion of stochastic and deterministic cointegration. Given the sample size used in this paper, the LS estimates of the cointegrating vector are likely to be substantially biased (see Banerjee, Dolado, Hendry and Smith 1986). Moreover, simple LS estimation does not allow hypothesis testing to be carried out on the estimated parameters of the cointegrating vector. For these reasons, LS estimation of the cointegrating regressions are carried out only for the purposes of testing the null hypothesis of no cointegration.

(1992) test for stationarity to a cointegrating framework. The test results presented in Table 3 (p. 16) suggest that for GC we are unable to reject the null hypothesis of cointegration for any of the tests at even the 10 per cent level, whereas for GI we are able to reject the null of cointegration at the 1 per cent level. These results suggest that C and GC are cointegrated, whereas C and GI are not.

In sum, the results for government consumption of goods and services lead us to conclude that the data are not sufficiently informative to distinguish between the presence or absence of cointegration. Therefore, in the next section we estimate the relationship between private consumption and government consumption of goods and services under both cointegration and no-cointegration assumptions. In marked contrast, the evidence corresponding to government investment allows us to strongly reject the hypothesis that C and GI are cointegrated. Thus, for government investment expenditure we estimate the extent of the relationship only under the maintained assumption of no-cointegration.

4. EMPIRICAL RESULTS

The first part of this section reports our estimates for the relationship between C and GC under the assumption of cointegration (that is, equation 12), while the latter presents the extent of the relationship estimated under the maintained assumption of no cointegration for both GC and GI (that is, equation 18). We note that in practice the residuals from equations (12) and (18), ε_t and v_t, will likely be serially correlated for a variety of reasons. Hence, we estimate both equations using Hansen's (1982) generalized method of moments (GMM). It should be stressed that estimates based on equations (12) and (18) are average effects over the sample period. That is, unless the composition of both private and government spending remains stable, the permanent income approach to government expenditures does not assume that the relationship between private and public consumption is stable over any particular subperiod.

Under the assumption of cointegration we estimate an empirical version of equation (12), that is,

$$c_t = \lambda_0 + \lambda_1 gc_t + \lambda_2 \Delta c_t + \lambda_3 \Delta gc_t + \varepsilon_t \tag{15}$$

using the regressors of equation (11) and their lagged values as instruments. Although this suggests a virtually infinite set of possible instruments, Monte Carlo evidence from Tauchen (1986) and Kocherlakota (1990) suggests that we should be parsimonious in the selection of our instruments. With this in mind, we estimate (20) and, later, equation (21) using rather parsimonious instrument sets.

Table 4 (p. 16) reports the estimates of λ_1, the implied value of α, the instrument sets and the J-tests of the overidentifying restrictions. The first two rows of Table 4 correspond to results generated using instruments that begin at time *t-1*. These results admit implied values of α about 0.5 and no evidence to reject the overidentifying restrictions. Thus, it appears that we find a significant degree of substitution between government and private consumption - a finding consistent with Kormendi (1983), Aschauer (1985) and Ahmed (1986). We should note, however, that our estimated degree of substitutability is somewhat larger than those for other countries. For instance, Aschauer (1985) and Ahmed (1986) estimate the degree of substitution between government and private consumption to be about 0.23 and 0.40 for the United States and the United Kingdom, respectively. Our larger response of private consumption to government spending may simply reflect differences across countries or the fact that these authors did not incorporate the distinction between cointegration and no cointegration into their work.

Although the results from the first two rows of Table 4 provides evidence for the substitution hypothesis, it is important to note that there may be a time aggregation problem if an agent's decision interval is finer than the data-sampling interval. This problem implies that an instrument at time *t-1* will not be orthogonal to the regression error and, hence, our previous estimates may not be valid. For our quarterly data the most recent admissible instruments are those measured at time *t-2*. The latter two rows of Table 4 present the results corresponding to instruments that begin at *t-2*. In contrast to the previous estimates, we now find implied values of α to range between -0.2 and -0.5, suggesting that private consumption responds positively to change in government spending. Again, none of the overidentifying restrictions are rejected. Thus, it appears that once we control for the possible effects of time aggregation, we find evidence consistent with that of Devereux, Head and Lapham (1994) and Karras (1994).[5]

We now estimate an empirical version of equation (18) under the maintained assumption of no cointegration, *viz.*,

$$\Delta c_t = \alpha_0 + \alpha_1 \Delta g_t + v_t \tag{16}$$

with g representing either *gc* or *gi*. Table 5 (p. 17) reports the estimates of α_1, the instrument sets and the J-tests of the overidentifying restrictions. Looking first at government investment expenditure we see that, except for one case, the point estimates of α_1 are not statistically significant from zero. This result is consistent with Kormendi (1983), who finds that public investment spending has little discernible effect on private consumption. However, the estimates of α_1 suggest that consumption and government investment are complements rather than substitutes regardless of instrument set. This conclusion that private consumption responds positively to changes in government investment is broadly consistent with Aschauer's (1989) result that increases in government capital lead to increases in the economy's productive capacity.

The latter rows of Table 5 present the estimation results for government

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5. It should be noted that Karras (1994) does not find a statistically significant relationship between the two types of consumption for Canada. The difference in conclusions may be attributed to the different empirical approach and/or different data (that is, Karras uses aggregate private and public consumption data).

consumption of goods and services. We obtain significant estimates of α_1 that range from 0.26 to 0.20 and find no evidence that the tests of the overidentifying restrictions can be rejected. Furthermore, unlike the cointegration case, the instrument set does not appear to have a large effect on our results. Under the no-cointegration assumption, we find a degree of substitutability between C and GC in line with the estimates found by Aschauer (0.23) for the United States. Hence, it appears that our distinction between cointegration and no cointegration is important for this type of analysis.

In a recent paper, Graham (1993) argues that an equation such as (21) is misspecified, since it omits the effects of real disposable income. This is motivated by Campbell and Mankiw's (1990) conclusion that consumption tracks disposable income too closely to be consistent with the permanent-income model. Indeed, if we simply regress the log of the first-difference of per capita disposable income on Δc, we find a parameter estimate of about 0.16 that is statistically significant at the 1 per cent level. To investigate Graham's claim for Canadian postwar data, we augment the government consumption of goods and services version of equation (21) with a disposable income term; that is, we estimate

$$\Delta c_t = \alpha_0 + \alpha_1 \Delta gc_t + \alpha_2 \Delta yd_t + \varepsilon_t \tag{17}$$

where yd_t is real per capita disposable income (see the Data Appendix for the source and a description). The results are reported in Table 6 (p. 17). It is readily apparent that the conclusions do not change when we include real disposable income, since the estimates of α_1 are not qualitatively different from those of equation (21). As well, we find that the parameter estimate for yd_t is never significant at even the 10 per cent level. This provides evidence in favour of the permanent-income model and suggests that Graham's claim may not be robust across different countries.

5. CONCLUDING REMARKS

This paper has examined whether or not there is a relationship between government and private consumption using Canadian data. We derive estimating equations between the two types of consumption under both cointegration and no-cointegration assumptions, an important distinction largely ignored in previous research. Our empirical results suggest that this distinction is crucial. For government consumption of goods and services our pretests do not allow us to firmly distinguish between the presence or absence of cointegration. Therefore, we estimate the relationship between private and government consumption under both cointegration and no-cointegration hypotheses. Under the null of cointegration and with instruments that control for the effects of time aggregation, private consumption and government purchases of goods and services appear to be complements. In contrast, under the no-cointegration null we find a statistically significant degree of substitution between private consumption and government consumption of goods and services that is robust to the inclusion of disposable income. For government investment, we are unable to find any evidence consistent with cointegration. We therefore estimate the response of private consumption under the no-cointegration hypothesis. From this data we are unable to find a statistically significant response. Economically, however, the parameter estimates suggest that private consumption and government investment are complements rather than substitutes.

In sum, our results suggest the presence of a relationship between government and private consumption. Unfortunately, we are unable to pin down the nature of this relationship using our approach. Given our results, the continuing debate on the response of private consumption to government spending shocks and the fiscal consolidation currently taking place in many industrialized countries, further research is needed in order to uncover the true impact of government spending on private consumption.

Data Appendix

This appendix describes the variables used in the study. All series were drawn from the CANSIM data base except the population series. The data definitions and reference numbers (provided in parentheses) are as follows: real consumption of non-durables and services (D20488-D20492-D20491-D29490), real government expenditures on goods and services (D20465) and real government investment (D20466). The total population 15 years of age and over is taken from the Labour Force Survey (LFSU1). Finally, real per capita disposable income (D20111) is deflated by the GDP implicit price deflator (D20011/ D20463) and by LFSU1.

TABLES:

Table 1:
Unit-Root Tests
Augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and Stock Tests[a]
Sample 1953Q1 to 1993Q2

Variable	ADF Lags	ADF t-statistic	PP Z_α-statistic	Stock's MZ_α-statistic
Consumption	2	-1.38	-2.62	0.53
Goods and services	5	-1.00	-8.05	0.53
Investment	0	-2.45	-5.96	0.54

a. Henceforth, "***", "**", "*" indicate significance at the 1, 5 and 10 per cent levels, respectively. The ADF critical values are calculated from MacKinnon (1991) while the PP critical values are taken from Fuller (1976). All test regressions include a trend term. For the ADF test we use the lag length selection procedure advocated by Ng and Perron (1994) and a 5 per cent critical value. The initial number of AR lags is set to the seasonal frequency plus 1 or 5. For the PP test, the long-run variance is estimated using a VAR prewhitened quadratic kernel estimator with a plug-in automatic bandwidth parameter as suggested by Andrews and Monahan (1992). For Stock's test we estimate the spectral density using an AR(4) spectral estimator.

Table 2:
Tests for the Null of No Cointegration
Augmented Engle-Granger (AEG), Phillips-Ouliaris (PO) and Stock Tests[a]

	Deterministic Cointegration		Stochastic Cointegration	
Tests	*gc*	*gi*	*gc*	*gi*
AEG t-statistic	-3.20* (1)	-0.96 (0)	-2.95 (2)	-1.34 (2)
PO Z_α-statistic	-13.53	-2.28	-8.29	-2.17
Stock's MZ_α-statistic	-9.67	-2.60	-19.76	-12.57

a. AEG critical values are calculated from MacKinnon (1991). PO Z_α and Stock's MZ_α critical values are taken from Haug (1992). The number in parentheses is the order of the AEG lags. For the AEG test we use the lag length selection procedure advocated by Ng and Perron (1994) and a 5 per cent critical value. The initial number of AR lags is set to the seasonal frequency plus 1 or 5. For the PO test, the long-run variance is estimated using a VAR prewhitened quadratic kernel estimator with a plug-in automatic bandwidth parameter as suggested by Andrews and Monahan (1992). For Stock's test we estimate the spectral density using an AR(4) spectral estimator.

Table 3:
Tests for the Null of Cointegration
Hansen and Shin Tests[a]

Government Spending	LC-statistic	MeanF-statistic	SupF-statistic	Shin η-statistic
gc	0.28	3.57	8.39	0.10
gi	2.26***	21.16***	64.04***	2.49***

a. The tests are based on VAR(1) prewhitened Phillips and Hansen (1990) parameter estimates.

Table 4:
Government Consumption of Goods and Services
Estimation of Equation (20) Under the Assumption of Cointegration[a]

Estimate of λ_1 (Standard Error)	Implied α	J-Test	Instrument Set
0.961*** (0.117)	0.471	0.238	Δc_{t-i}, Δgc_{t-i} (i = 1,2)
0.954*** (0.097)	0.537	1.322	Δc_{t-i}, Δgc_{t-i} (i = 1,2,3)
0.840*** (0.154)	-0.543	0.573	Δc_{t-i}, Δgc_{t-i} (i = 2,3)
0.773*** (0.162)	-0.201	0.108	Δc_{t-i}, Δgc_{t-i} (i = 2,3,4)

a. Henceforth, GMM estimation is performed with the Bartlett kernel and the truncation parameter set equal to one. All instrument sets used in the current analysis include a constant.

Table 5:
Estimation of Equation (21) Under the Assumption of No Cointegration

Government Spending	Estimate of α_1 (standard Error)	J-Test	Instrument Set
Investment	0.123 (0.138)	4.461	Δc_{t-i}, Δgi_{t-i} (i = 1,2)
	0.205* (0.128)	4.290	Δc_{t-i}, Δgi_{t-i} (i = 1,2,3)
	0.354 (0.220)	2.209	Δc_{t-i}, Δgi_{t-i} (i = 2,3)
	0.341 (0.243)	1.944	Δc_{t-i}, Δgi_{t-i} (i = 2,3,4)
Goods and Services	-0.200* (0.117)	3.709	Δc_{t-i}, Δgc_{t-i} (i = 1,2)
	-0.213* (0.118)	7.801	Δc_{t-i}, Δgc_{t-i} (i = 1,2,3)
	-0.263* (0.145)	5.740	Δc_{t-i}, Δgc_{t-i} (i = 2,3)
	-0.156 (0.118)	9.212	Δc_{t-i}, Δgc_{t-i} (i = 2,3,4)

Table 6:
Government Consumption of Goods and Services
Estimation of Equation (22): The Disposable Income Augmented Equation

$\hat{\alpha}_1$ (standard error)	$\hat{\alpha}_2$ (standard error)	J-Test	Instrument Set
-0.255* (0.148)	-0.226 (0.224)	3.109	Δc_{t-i}, Δgc_{t-i}, Δy_{t-i} (i = 1,2)
-0.194* (0.120)	0.053 (0.156)	9.273	Δc_{t-i}, Δgc_{t-i}, Δy_{t-i} (i = 1,2,3)
-0.224 (0.154)	0.277 (0.232)	7.751	Δc_{t-i}, Δgc_{t-i}, Δy_{t-i} (i = 2,3)
-0.119 (0.113)	0.032 (0.234)	11.811	Δc_{t-i}, Δgc_{t-i}, Δy_{t-i} (i = 2,3,4)

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