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INTERTEMPORAL ANALYSIS OF STATE AND LOCAL GOVERNMENT SPENDING: THEORY AND TESTS

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ABSTRACT

Do state and local governments smooth their consumption spending across years, or is their spending driven mainly by contemporaneous changes in resources? We design a test to determine which view of state and local spending is more consistent with the data. We find that state and local spending is determined primarily by current (as opposed to permanent) resources. That is, despite their apparent ability to skirt balanced budget laws, states and localities do not typically smooth their expenditures over time.

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1. INTRODUCTION

One of the distinguishing features of modern economics is a focus on intertemporal considerations. Much attention is devoted to testing the theoretical prediction that the behavior of an agent today depends upon his expectations about the future economic environment. For years, discussions of both household consumption decisions and business investment decisions have been dominated by controversy over the extent to which they can be explained by models focusing on intertemporal decision-making.[1]

In contrast, the analysis of another important component of aggregate demand, state and local government spending, still typically ignores such issues. Both theoretical and empirical analyses of state and local government behavior generally *assume* that all spending during a given period depends only on resources available in that period.[2] This absence of attention to the dynamics of aggregate state and local spending is particularly striking given its role in determining aggregate demand--state and local purchases of goods and services were $642 billion in 1991, about 11 percent of GDP, and about $90 billion more than that year's nonresidential investment expenditures.[3]

While one might argue that intertemporal decision-making models are obviously inappropriate for understanding government behavior, such models have been employed with some success to analyze the fiscal behavior of the federal government. (See, for example, Barro [2], Mankiw [12], or Poterba and Rotemberg [14].) However, a factor that must be considered in this context is that state and local governments, unlike the federal government, face balanced-budget laws. Do such laws make it impossible for these governments to borrow so as to smooth consumption flows over time? Not necessarily. Inman [10], Leonard [11] and others have emphasized that these laws require only that a balanced budget be submitted prior to the fiscal year, not that spending actually adhere to the budget. As Leonard notes, "Even states with such limits can experience deficits if ... there is more of a revenue shortfall than officials were able to foresee. Because most balance rules require equating an estimate of revenues with an estimate of expenditures, there is considerable slack in the budget balance requirement of most state budget processes" (p. 224). For example, in 1992, Governor Jim Florio of New Jersey submitted a balanced budget to the state legislature. However, the *New York Times* noted that "balance" was obtained on the basis of blatantly unrealistic assumptions: --"[The Governor] hopes that an economic recovery will yield higher state tax collections, that Washington will provide a windfall of Medicaid reimbursements and that the Legislature will approve revised accounting for state pension funds" (January 29, 1992, p. A20).

In short, states and localities can and do circumvent balanced budget rules. As a consequence, there are no effective *legal* restrictions preventing sub-federal governments from smoothing spending over time. Of course, just as for households, capital market imperfections and myopia may reduce the amount of such smoothing. However, there is at least some evidence that sub-federal governments can offset revenue shocks with borrowing. An example that received substantial publicity was during 1988, when the Dukakis administration in Massachusetts borrowed $200 million to cover a short-run revenue shortfall.[4]

Another possible objection is that even if their spending is not legally tied to current resources, it is implausible that state and local decision-makers act as if they are maximizing an intertemporal utility function. Several investigators such as Inman [10] have suggested that public sector decision-making is backward-looking, as opposed to the forward-looking view embodied in intertemporal utility maximization. On the other hand, our conversations with several state and local officials have indicated that their time horizons are longer than just a single budget year. Indeed, some jurisdictions have "rainy day funds" which are designed to allow them to smooth spending over time, and others operate on multi-year budget cycles. Of course, none of this proves that an intertemporal perspective is correct for state and local governments; it only suggests that such a model is worth testing.

The purpose of this paper is to investigate the extent to which state and local spending can be rationalized by a model in which decisions are based on the permanent (as opposed to current) level of resources available to that sector. We specify a model in which state and local government spending is generated by the maximization under uncertainty of an intertemporal utility function, and test this model with annual time series data from 1934 to 1991. The model is presented in section 2. Section 3 discusses the data and the econometric strategy. The results are presented in Section 4. The major finding is that essentially all of current spending is determined only by contemporaneous variables--despite their ability to skirt balanced budget laws, states and localities do not typically smooth their expenditures over time. Section 5 concludes with a summary and a discussion of some implications of the results.

2. THE MODEL

Intertemporal approaches to decision-making assume that agents are forward-looking. They make their decisions regarding today's expenditures on the basis of their expectations about future resources. To apply this notion to state and local government decision-making, we assume the existence of a decision-maker who is concerned about the flow of government services not only in the present year but into the indefinite future as well. This decision- maker's goal is to maximize the expected present value of utility (which depends on government expenditures) subject to an intertemporal budget constraint. Beginning with Hall [6], tests of this kind of model have adopted the following strategy. First, solve the decision-maker's maximization problem using dynamic programming methods. Second, note any restrictions that the solution places on the lag distribution of the choice variable (spending). Third, analyze the time series data to determine whether or not these restrictions are violated. To the extent that the data are consistent with the theoretical restrictions, it suggests the presence of rational, forward-looking planning. Alternatively, if the restrictions are violated, then myopia, short-run constraints, or backward-looking behavior may be present.

Applying this procedure to our problem, we initially assume that the goal is to maximize the expected discounted value of a utility function that depends upon the flow of government services (as measured by current expenditure).[5] Letting $U(\cdot)$ denote the period-specific utility function, the goal is to maximize

$$V_t = E_t \{\sum_{s=0}^{\infty} \beta^s U(G_{t+s})\} , \tag{2.1}$$

where E_t denotes expectations taken using information available through the end of period t, $\beta = 1/(1+\delta)$ and δ is the pure rate of time preference, and G_t is the level of state-local government spending on nondurable goods and services in period t. We assume that such state-local spending is the only argument in the utility function for clarity of presentation alone. We explore below the consequences of permitting durable state and local goods, spending on transfer payments, and other types of consumption to affect utility. An attractive feature of this approach is that it does not require us to specify whose preferences are represented by $U(\cdot)$. For example, the utility function might be that of a representative resident, or it might depict the preferences of a bureaucrat whose utility depends on the size of his budget. As long as the decision-maker's preferences are stable, equation (2.1) serves as a useful framework. This is, of course, analogous to the usual assumption of a stable preference ordering for the representative agent in the consumption function literature.

The state-local sector is constrained by the intertemporal budget constraint that initial wealth plus the present value of resources at least cover the present value of expenditures. In any time period, resources (R_t) are the sum of own-source revenues plus outside grants-in-aid from the federal government.[6] Denoting end of period net wealth by W_t, state-local governments must satisfy the present value budget constraint:

$$W_{t-1} + \sum_{s=0}^{\infty} \psi^s \left(R_{t+s} - G_{t+s} \right) = 0 \tag{2.2}$$

where $\psi = 1/(1+r)$ and r is the constant real rate of interest.

The optimal spending path is characterized by the system of Euler equations

$$E_t \left\{ [U'(G_{t+s})/U'(G_{t+s-1})] - [(1+\delta)/(1+r)] \right\} = 0, \quad s = 1,...,\infty. \tag{2.3}$$

In words, the marginal rate of substitution between government consumption in adjacent periods is equated (in expected value) to the intertemporal relative prices--the ratio of one plus the rate of time preference to one plus the real interest rate. Taking natural logarithms, dropping the expectations, and adding an expectational error yields the *ex post* relationship:

$$\ln U'(G_t) - \ln U'(G_{t-1}) = \ln[(1+\delta)/(1+r)] + v_t \; . \tag{2.4}$$

Equation (2.4) is Hall's [6] result that the marginal utility of consumption (G_t) is a martingale. Moreover, if expectations are formed rationally:

$$E_{t-1}\{v_t\} = 0 \; . \tag{2.5}$$

Equation (2.5) indicates that with rational expectations, the expectational error is uncorrelated with any information available at the time the decision is made; i.e. any variable observed through the end of period t-1.[7]

In summary, the model assumes that: i) government spending decisions are the outcome of maximizing the discounted value of a time-separable objective function; ii) there are no credit market constraints on the state and local government sector; and iii) expectations are formed rationally. For brevity, and in analogy to the modern literature on private consumption, henceforth we will refer to this as the "permanent income model." In contrast, spending that is determined only by current resources (due to myopia, credit market constraints, etc.) will be characterized as "Keynesian", in analogy to the Keynesian consumption function in the macroeconomics literature.[8]

In practice, the permanent income model can be tested by estimating an equation of the form

$$\Delta \ln G_t = \alpha_0 + \sum_{i=1}^{k} \alpha_i \Delta \ln G_{t-i} + \eta_t \,, \tag{2.6}$$

where the α's are parameters to be estimated and η_t is a random error term. If the marginal utility of G_t is a martingale, then $\alpha_i = 0$ for all i. Note that if the decision-maker's utility function displays constant relative risk aversion, then equation (2.6) is exact. Otherwise, it should be viewed as a first-order logarithmic approximation. Both logarithmic and linear specifications appear in the literature; for our purposes the logarithmic approach is more convenient.

A finding that the α's are significantly different from zero rejects the permanent income model. However, the mere fact that the α's are not zero does not tell us the quantitative significance of the rejection. It is possible, for example, that 90 percent of spending is determined in accordance with the permanent income model. If so, it is only the remaining 10 percent of spending, due to Keynesian spenders, that produces the rejection. Recently, Campbell and Mankiw [4] (referred to hereafter as C-M) have suggested an alternative econometric test that allows one to measure the quantitative significance of departures from the model. As already noted, under the null hypothesis that state-local government spending is determined by a permanent income model, changes in G_t are due only to changes in the permanent resources of the governments. Letting ε_t be the percentage change in permanent resources, we have

$$\Delta \ln G_t = \upsilon + \varepsilon_t \ . \tag{2.7}$$

where υ is the *ex ante* planned growth rate of G_t. In contrast, in some instances changes in G_t may be determined by Keynesian decision-makers--whenever current resources go up by a certain percentage, so does current spending, and vice versa. That is,

$$\Delta \ln G_t = \Delta \ln R_t \ , \tag{2.8}$$

where R_t denotes the resources available to the state-local sector. Now, suppose that λ is the fraction of state and local government spending determined by current resources, so that $(1\text{-}\lambda)$ is the proportion of spending consistent with the permanent income hypothesis. In other words, λ represents the dollar weighted proportion of governments that are Keynesian spenders. The actual change in government spending is just a weighted average of (2.7) and (2.8):

$$\Delta \ln G_t = (1-\lambda)\upsilon + \lambda \Delta \ln R_t + (1-\lambda)\varepsilon_t \ . \tag{2.9}$$

Equation (2.9) is the centerpiece of the empirical analysis below. Our goal is to estimate λ, the fraction of state-local spending determined by current resources. An empirical estimate of $\lambda=0$ indicates that state-local government consumption is determined in accordance with the permanent income model. At the other extreme, an estimate of $\lambda=1$ indicates that this important component of aggregate demand is determined only by the current resource flows to the state-local sector. Values of λ in between tell us the proportion of state-local spending that is Keynesian.

The appealing simplicity of equation (2.9) is a direct result of the stringent assumptions adopted above. In particular, its derivation assumed that the real interest rate is constant, and that the marginal utility of state and local government spending on nondurable goods and services is affected by neither the level of state and local spending on durables, the level of state and local spending on transfer payments, the level of private consumption, nor the level of federal government spending. When fluctuations in real interest rates are introduced, direct inspection of equation (2.4) indicates that changes in G_t will be related to changes in r_t. In the same way, if state and local government spending is not separable in the utility function, then ΔlnG_t will also depend upon changes in the levels of state and local durable goods, state and local transfer payments, private consumption, and consumption of federal government services. We can incorporate these considerations by letting D_t denote state-local durable spending, T_t state and local transfer payments, C_t private consumption, and F_t federal government consumption, and then re-writing (2.9) as

$$\Delta \ln G_t = (1-\lambda)\upsilon + \lambda\Delta\ln R_t + \theta_1 r_t + \theta_2\Delta\ln D_t + \theta_3\Delta\ln T_t + \theta_4\Delta\ln C_t + \theta_5\Delta\ln F_t + (1-\lambda)\varepsilon_t \ . \qquad (2.10)$$

As C-M note, this specification is exact if the period-specific utility function, $U(\bullet)$, is Cobb-Douglas. Otherwise, it may again be viewed as a log-linear approximation to an arbitrary utility function. If (2.10) is the appropriate specification, then incorrect inferences regarding λ may be made by estimating (2.9).

A problem that arises in the estimation of both (2.9) and (2.10) is that simultaneity bias may emerge if ordinary least squares (OLS) is used. Consider equation (2.9). To the extent that changes in current resources are not perfectly anticipated, they will lead to revisions in permanent resources. Thus, ΔlnR_t and ε_t will be correlated and OLS estimates will be biased; the direction of the bias may be positive or negative. (Similar arguments apply to equation (2.10).) An alternative estimation strategy is discussed below.

3. DATA AND ECONOMETRIC ISSUES

3.1 Data

We estimate the model using annual National Income and Product Accounts (NIPA) data from 1934 to 1991. To test the sensitivity of our results to the inclusion of the Great Depression and World War II, we also estimate the model using the 1945-1991 subsample. All dollar figures are converted to 1982 terms using the state and local expenditure deflator, and then put on a per capita basis by dividing by population. State and local spending, G_t, is the sum of nondurable goods and services. In principle, G_t should also include the services generated by the stock of durables and structures owned by state and local governments. However, creating such a time series requires a number of arbitrary assumptions,[9] so, following Hall [6] and others, we decided to omit them from the analysis. Implicitly, this amounts to assuming that the services generated by durables enter separably into the utility function, an assumption that is tested below.

Resources available to the state and local sector, R_t, is the sum of personal tax and nontax receipts, corporate profits tax accruals, indirect business tax and nontax accruals, and federal grants-in-aid. D_t is the level of state and local durable goods purchases, T_t is state and local transfer payments, C_t is personal consumption of nondurable goods and services, and F_t is federal government expenditures on nondurable goods and services. The real interest rate (r_t) is computed by taking the nominal interest rate and subtracting from it the *ex post* rate of change of prices for state and local goods and services.[10] The nominal interest rate is the yield on Baa corporate bonds as computed by Moody's Investor's Service.[11] The inflation rate is the annual rate of change of the state and local price deflator during the year. In principle, another possible source of intertemporal price variation is changes in matching rates for federal grants. In practice, however, most matching grants apply to expenditures on either durables or welfare, and these are excluded from our measure of G_t.[12]

Table 1 shows the mean values of the logarithms of G_t, R_t, C_t, D_t, T_t and F_t, the mean values of the first differences of the logarithms, and the associated standard errors. It is particularly noteworthy that although on average real state and local spending and resources have been increasing over time (the means of ΔlnG_t and ΔlnR_t are positive), there has been substantial volatility in both these growth rates, as indicated by the relatively large standard errors.

3.2 ECONOMETRIC ISSUES

As indicated in Section 2, the basic estimating equation is:

$$\Delta \ln G_t = (1-\lambda)\upsilon + \lambda \Delta \ln R_t + (1-\lambda)\varepsilon_t \tag{3.1}$$

where υ and λ are parameters, and ε_t is a white noise error. If one cannot reject the hypothesis that $\lambda=0$, then the data are consistent with the permanent income model; similarly, if one cannot reject the hypothesis that $\lambda=1$, then the data are consistent with the Keynesian model. The major econometric problems arise because ε_t, the revision in the decision-maker's forecast of future resources, is likely to be correlated with ΔlnR_t, the change in current resources. Hence, OLS estimation is inappropriate. Following C-M's procedure, we use an instrumental variables (IV) estimator, and employ lagged values of various variables as instruments.

As C-M note, a nice feature of the instrumental variables approach is that it provides the basis for a test of some over-identifying restrictions implied by equation (3.1). Suppose that the set of instruments is $X_{1t}, X_{2t}, \ldots X_{kt}$. Consider the system:

$$\Delta \ln G_t = \pi_0 + \pi_1 X_{1t} + \pi_2 X_{2t} + \ldots + \pi_k X_{kt} + \eta_{Gt} \tag{3.2a}$$

$$\Delta \ln R_t = \gamma_0 + \gamma_1 X_{1t} + \gamma_2 X_{2t} + \ldots + \gamma_k X_{kt} + \eta_{Rt} \ . \tag{3.2b}$$

When the instruments are variables dated $t\text{-}1$ and earlier, equation (3.2a) is a conventional test of the restrictions implied by equations (2.4) and (2.5). The data are consistent with the permanent income model if $\pi_1=\pi_2=\ldots=\pi_k=0$. Equation (3.2b) is just the first stage equation in the IV estimation of (3.1). The key thing to note is that equation (3.1) places over-identifying restrictions on the π's and γ's of equations (3.2a) and (3.2b). To see why, consider a one unit increase in X_{1t}. According to (3.2), this raises ΔlnG_t by π_1, and ΔlnR_t by γ_1. However, equation (3.1) says that whenever ΔlnR_t increases by a given amount, then ΔlnG_t must increase by λ times that amount. Hence, $\pi_1 = \lambda\gamma_1$. Similarly, $\pi_2=\lambda\gamma_2$, $\pi_3=\lambda\gamma_3$, ..., $\pi_k=\lambda\gamma_k$. In effect, testing these restrictions amounts to testing the assumptions behind equation (3.1), *viz.*; the functional form, the appropriateness of aggregating over jurisdictions, the exclusion of other possible right hand side variables, and the orthogonality of the error term.

These restrictions can be tested by taking the residuals from the IV estimates of (3.1) and regressing them on the instruments.[13] The relevant test statistic is T times the R^2 of this regression, where T is the sample size. It is distributed as a chi-square with $(k\text{-}1)$ degrees of freedom.

The issue now is the selection of instruments, i.e., the X's of equations (3.2). Following common practice, we employ lagged values of ΔlnG_t as well as the lagged values of the right-hand side variables in equation (2.10). The question then becomes which lags to use. First lags should not be employed for three reasons: i) Omitting first lags is likely to reduce problems associated with time aggregation; ii) If some nondurable goods in the NIPA data are actually durable, then ε_t may have a moving average structure and first lags of the variables will be correlated with ε_t; and iii) In the presence of transitory spending shocks, ε_t again has a moving average component. In light of these considerations, we exclude first lags from the list of instruments. However, all lags dated *t-2* and earlier are candidates for instruments. Normally, in a finite sample the only cost of "too many" instruments is a loss of efficiency in the parameter estimates. Here, however, the use of too many lags may also reduce the power of our restriction tests. Intuitively, lags that do not "really" belong will have zero coefficients in both equations (3.2a) and (3.2b), tending to bias toward acceptance the test of the hypothesis that $\pi_i=\lambda\gamma_i$ for all values of i. To provide the most stringent test possible, we select the most parsimonious set of instruments consistent with the data. To do so, we estimate equation (3.2b) with lags 2 through 5 of the variables, and then with successively shorter lag lengths. Each time that the lag length is reduced, we test for the appropriateness of the shorter lag length using a likelihood ratio test. In this way, the set of instruments is pared down as far as the data permit.

A final econometric issue is heteroskedasticity of the error term. As noted earlier, the error term in equation (3.1) may be a first-order moving average, in which case standard OLS and IV estimation will produce inconsistent standard errors. The estimates reported below therefore use White's [16] covariance matrix estimator to provide consistent estimates of the standard errors in the presence of conditional heteroskedasticity.

4. RESULTS

To begin, we estimated (3.1) by OLS. The outcome, reported in column 1 of Table 2, indicates a value of λ of 0.815 with a standard error of 0.0902. On this basis, one would reject the hypothesis that aggregate state and local spending follows the permanent income model ($\lambda=0$), but not reject the hypothesis that 100 percent of the dollars are spent by Keynesian decision-makers ($\lambda=1$).

As emphasized in section 2, there is good reason to suspect that the simple model is misspecified. In particular, an equation like (3.1) might falsely reject the permanent income model if the real interest rate is not constant, or if there are nonseparabilities between state and local spending and other variables in the utility function. We therefore estimate equation (2.10), which allows for these possibilities. The OLS results are reported in column (2). The impled value of λ is essentially unchanged.

As stressed in section 3, OLS may be an inappropriate estimation method, because the right-hand side variables are potentially endogenous. In column (3), therefore, we present the results when equation (2.10) is estimated by instrumental variables.[14] The value of λ increases to 0.945 (s.e.=0.133). One still may reject the hypothesis that $\lambda=0$, but one may not reject the hypothesis that $\lambda=1$.

However, there is no reason to put much faith in this result if the over-identifying restrictions implied by the model are violated. The chi-square test statistic associated with these over-identifying restrictions is 13.89 with eight degrees of freedom. The critical value of the chi-square distribution at the five percent significance level is 15.5. Hence, the overidentifying restrictions are not rejected by the data.

It is informative to compare the OLS and IV estimates of λ. The direction of bias in the OLS estimate depends on the correlation between the current change in resources (ΔlnR_t) and the innovation in permanent resources (ε_t). The fact that λ rises with IV estimation indicates that this correlation is negative. This is consistent with the notion that state and local decision-makers believe that policies that result in unanticipated increases in current resources may decrease permanent resources. One possible reason for this phenomenon is that swings in current tax rates in response to unanticipated economic or demographic shocks may lead to the loss of future tax base. For example, if states and cities raise tax rates (and revenues) during "hard times," firms and individuals may move to other jurisdictions in the future.[15] To determine the validity of this conjecture would require estimating a structural model of permanent resource determination, a task that is beyond our scope.

Another observation suggested by the estimates in column (3) is that on a one-by-one basis, neither r_t, $\Delta \ln D_t$, $\Delta \ln T_t$, $\Delta \ln C_t$, nor $\Delta \ln F_t$ is statistically significant. Indeed, one cannot reject the joint hypothesis that $\theta_1=\theta_2=\theta_3=\theta_4=\theta_5=0$ (the significance level of the test is 0.206). We therefore performed two additional exercises to assess the robustness of the estimate of λ. In the first, we estimated λ by instrumental variables, imposing the constraint that θ_1 through θ_5 equal zero. (See column (4)). Second, we deleted various combinations of right-hand side variables to see if, with a more parsimonious set of regressors, it was possible to obtain statistically significant estimates of at least some of the θ's. (See column (5)).[16] As the results in the table indicate, our fundamental finding does not change--one cannot reject the hypothesis that $\lambda=1$.[17] Essentially 100 percent of changes in state and local governments spending on nondurable items are driven by changes in contemporaneous resources.

One final matter of possible concern relates to our sample period, which includes part of the Great Depression and World War II. Are these extraordinary years dominating the results? To address this question, we re-estimated the models with data only from 1945 to 1991. Using the postwar sample, IV estimation of the model with the entire set of right-hand side variables (corresponding to column (3) in Table 2) leads to a λ of 1.29 (s.e. = 0.418). IV estimation of the simple model (corresponding to column (5) in Table 2) leads to a λ of 1.01 (s.e. = 0.277). As expected, with a smaller sample, the estimated standard errors are larger. Nevertheless, the basic message is the same--state and local nondurable spending follows a Keynesian rather than a permanent income pattern.[18]

5. CONCLUSION

In this paper we have examined the extent to which intertemporal considerations play a role in determining spending flows by state and local governments. Taking advantage of a procedure recently suggested by Campbell and Mankiw [4], we found that essentially 100 percent of the growth rate of state and local spending on nondurable items is determined by growth in the decision-maker's contemporaneous level of resources.[19] This is in direct contrast to various studies which have shown that the federal government's behavior is characterized by a longer run decision-making horizon.

This finding has important policy implications. The federal government often seeks to influence the level of state and local spending by changing the resources available to that sector. This can be done directly via grants-in-aid, or indirectly by such measures as the deduction for sub-federal taxes on federal tax returns. Our results suggest that the effects of such measures on aggregate state and local spending on nondurable items are likely to be quite immediate because unlike consumers, state and local governments do not spread transitory changes in their incomes over time. This observation is particularly important if the measures are being undertaken as part of a macroeconomic stabilization program.

TABLE 1

SUMMARY STATISTICS[*]

Variable	Mean[a]
lnG_t State-Local Nondurable Goods and Services	6.76 (0.364)
lnR_t State-Local Resources	7.14 (0.327)
lnD_t State-Local Durable Goods	3.66 (0.591)
lnT_t State-Local Transfers	5.25 (0.500)
lnC_t Personal Consumption Nondurable Goods and Services	9.01 (0.134)
lnF_t Federal Nondurable Goods and Services	6.75 (0.450)
ΔlnG_t Growth Rate of G_t	0.0165 (0.0524)
ΔlnR_t Growth Rate of R_t	0.0138 (0.0357)
ΔlnD_t Growth Rate of D_t	0.0176 (0.179)
ΔlnT_t Growth Rate of T_t	0.0501 (0.123)
ΔlnC_t Growth Rate of C_t	0.00380 (0.0529)
ΔlnF_t Growth Rate of F_t	0.0317 (0.249)
r_t *Ex Post* Real Interest Rate	1.75 (4.05)

[*]All variables other than r_t are in real, per capita terms.

[a]Numbers in parentheses are standard errors.

TABLE 2

PARAMETER ESTIMATES*

Variable	(1) OLS	(2) OLS	(3) IV	(4) IV	(5) IV
ΔlnR_t	0.816	0.814	0.945	1.01	1.02
	(0.0904)	(0.100)	(0.133)	(0.206)	(0.180)
ΔlnD_t		-0.0418	-0.153		-0.137
		(0.0325)	(0.0868)		(0.0786)
ΔlnT_t		0.0264	0.0946		
		(0.0322)	(0.124)		
ΔlnC_t		-0.123	-0.108		
		(0.102)	(0.194)		
ΔlnF_t		-0.324	-0.0660		-0.0830
		(0.0222)	(0.0480)		(0.0361)
r_t		-0.000594	-0.000761		
		(0.00204)	(0.00231)		
Constant	0.00409	0.0109	0.0113	0.000997	0.00707
	(0.00628)	(0.0108)	(0.0135)	(0.00743)	(0.00890)

*All variables are defined in Table 1. Numbers in parentheses are standard errors.

Notes

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1. Hall [6] and Campbell and Mankiw [4] are two examples of the voluminous modern literature on the consumption function. Summers [15] and Abel and Blanchard [1] discuss intertemporal models of business investment.

2. For excellent surveys of this work, see Inman [9] and Mieszkowski and Zodrow [13].

3. *Economic Report of the President 1992*, pp. 298-99.

4. *New York Times*, August 22, 1988, p. B6.

5. Note the contrast to Barro's [2] model, in which the government's choice variable is the sequence of tax rates. While it would be interesting to analyze a model in which there is both expenditure and tax rate smoothing, it is beyond the scope of this paper. However, our econometric procedure (discussed below) does take into account the endogeneity of tax revenues.

6. In work using a dynamic framework, Holtz-Eakin and Rosen [7] found that one cannot reject the hypothesis that total resource flows determine construction spending; entering grants and own-source revenues separately does not significantly increase the explanatory power of the model.

7. Subsequent to Hall's research, several investigators have proposed that consumption may follow an ARMA(1,1) process, and, hence, equation (2.5) may not hold. As indicated in Section 3.2, below, our econometric procedure circumvents this problem.

8. We make no attempt to distinguish between myopia and credit market constraints. Flavin [5] tries to test for liquidity constraints on individuals by introducing the unemployment rate as a proxy variable for the presence of such constraints. While this approach is interesting, there does not seem to be any available variable that could serve as a compelling proxy for liquidity constraints faced by state and local governments.

9. See Boskin, Robinson, and Huber [3] and Hulten and Schwab [8] for careful discussions of problems in measuring the service flows from the state-local capital stock.

10. The use of *ex post* rather than *ex ante* values of the interest rate is common in such analyses; see, e.g., C-M. Under the rational expectations hypothesis, the difference between *ex post* and *ex ante* interest rates in period t is a function of new information in period t. This new information is embodied in the error term in period t. Hence, it is appropriate to use the *ex post* rate provided that it is instrumented with lagged interest rates, which are correlated with the current rate but uncorrelated with the error.

11. It might be preferable to use the rate of return on state and local bonds instead, but we have no constistent time series on this variable for the entire sample period. However, for the post 1950-period, for which the corporate and state-local series overlap, the correlation between the two is very high--the R^2 of a regression of the state-local rate on the corporate rate is 0.921. Hence, we doubt if any serious errors are introduced by employing the corporate rate.

12. Algebraically, let m_t be the federal matching rate in year t, let p_t be the acquisition price of state and local goods in period t, and let r_{nt} be the nominal interest rate in year t. Then the real interest rate, r_t, in period t is given by:

$$r_t = r_{nt} - \left(\frac{(1-m_{t+1})p_{t+1}}{(1-m_t)p_t} - 1 \right) .$$

In effect, we assume that $m_{t+1}=m_t$, which is sensible given our definition of G_t.

13. To see this, consider a simplified version of equations (3.1) and (3.2) in which we ignore the intercepts and only X_{1t} appears on the right side. In these circumstances, substituting equation (3.2b) into equation (3.1) yields

$$\Delta \ln G_t = \lambda(\gamma_1 X_{1t} + \eta_{Rt}) + (1-\lambda)\varepsilon_t . \qquad \text{(A)}$$

Adding and subtracting $\pi_1 X_{1t}$ to the right hand side, this may be written

$$\Delta \ln G_t = \pi_1 X_{1t} + (\lambda\gamma_1 - \pi_1)X_{1t} + \lambda\eta_{Rt} + (1-\lambda)\varepsilon_t . \qquad \text{(B)}$$

Calling the last three terms on the right hand side of equation (B) η_{Gt}, we have

$$\Delta \ln G_t = \pi_1 X_{1t} + \eta_{Gt} \qquad \text{(C)}$$

which corresponds to equation (3.2a). Notice from equation (B) that, for example, if the proportionality restriction $\pi_1 = \lambda\gamma_1$ is satisfied, the error term in (C), η_{Gt}, is uncorrelated with X_{1t}. Should, the restriction be violated, however, X_{1t} will be correlated with the error term, and this violation of the restrictions will be detected by a regression of the residuals on the instruments (in this case, X_{1t}). Violations stemming from other sources (such as omitted variables) will be detected in the same way.

14. The instruments, selected by the algorithm discussed above, are lags 2 and 3 of the various variables.

15. We are grateful to a referee for this suggestion.

16. The fact that personal consumption enters the equation insignificantly lends some support to our underlying assumption that the error term in equation (2.10) is a

consequence of expectational errors rather than taste shocks. In the presence of taste shocks, one would typically expect both public and private consumption to be affected. This implies that C_t would be correlated with that component of the error term reflecting the taste shock. Our results suggest that this is not the case. It is possible, of course, to imagine shocks that would affect the marginal utility of public spending while leaving the marginal utility of private spending unchanged. Thus, this result needs to be interpreted with some caution.

17. In the column (4) results, the test of the overidentifying restrictions yields a significance level of 0.131; for the column (5) results, the significance level is 0.185. Hence, as was the case above, one cannot reject the model's overidentifying restrictions.

18. We also estimated the model using quarterly data from 1960.1 to 1987.3. Such an exercise may be problematic if state and local governments do intrayear smoothing even in the presence of binding annual balanced budget constraints. On the other hand, use of quarterly data does allow us to examine a period of time which is sufficiently short that major regime shifts are unlikely. In any case, this exercise yields a somewhat lower point estimate of λ, but one which is still insignificantly different from one and statistically different from zero.

19. Of course, the fact that spending on nondurable items follows a Keynesian pattern does not mean that all spending behaves that way. For example, Holtz-Eakin and Rosen's [7] study of panel data on individual communities indicated that at least for some communities, the permanent income model well characterizes spending behavior on capital goods.

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