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by

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ABSTRACT

We construct a model that considers the *direct* effects, if any, of government spending on the attitudes of a typical consumer toward risk, time preference, and intertemporal substitution. The null hypothesis is that a growing government sector does not affect the consumer's behavior, and the alternative is that it causes him to become less risk averse, more impatient to consume now rather than in the future, and less responsive to changes in real interest rates. If the alternative hypothesis is correct, then government growth may lead to lower economic growth. Using Greek annual aggregate data, 1960-1990, we can reject the null hypothesis.

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1. Introduction

During the past three decades, the size of government, measured by the share of government expenditure in GDP, has been growing in many countries, notably those in the OECD. The consequences of this phenomenon have not been fully assessed, however, and the issue is still fervently debated among economists and policy-makers.

In this paper, we construct a model that considers the *direct* effects, if any, of government growth on the attitudes of a typical consumer toward risk, time preference, and intertemporal substitution in consumption. Our null hypothesis is that a growing government sector does not affect these attitudes, and the alternative is that it causes a typical consumer to become less risk averse, more impatient to consume now rather than in the future, and less responsive to changes in real interest rates. If the alternative hypothesis is correct, then government growth may cause economic growth to decline by causing private saving and market efficiency to decline. For if a consumer becomes less risk averse, then he will insure less against uncertain future income by saving less. And if he becomes less willing to postpone consumption, then saving will fall at every given level of the real interest rate. As a result, current and future real interest rates will rise, thus lowering investment in physical capital. Productivity and long-run economic growth will decline. To our knowledge, no formal model exists that examines these effects, although previous studies, e.g., Bean (1986), explicitly consider the effects of government expenditure on consumer behavior.

We use an intertemporal utility maximization model to address the issue. Following Bean (1986), we include government expenditure in the utility function but not in the budget constraint of a typical consumer, i.e., the model is not Ricardian.[1] Using Greek annual aggregate data, 1960-1990, we obtain reasonable parameter estimates that lead to the rejection of the null hypothesis. Our

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[1] Ricardian Equivalence could easily be incorporated into the model, e.g., as in Aschauer (1985). It does not alter out estimating equation, however, and thus our empirical results are not affected by its exclusion.

motivation to use Greek data comes from the spectacular growth of the size of government during the 1980s, which is widely believed to have caused the recent macroeconomic imbalances [see OECD (1991b)].[2] It also comes from the suggestion of the OECD (1987b, p. 38) that fast-growing government transfers may have caused a "rentier mentality" in the Greek society, and may have created a "climate of complacency, dissociating income from work effort, reducing work incentives and favoring consumption." We present the model in section 2 and the empirical results in section 3. Section 4 concludes the paper.

2. The Model

2.1 The Consumer's Problem

We consider a representative consumer who maximizes a multi-period utility function. Having taken all the relevant constraints into account and having formed his expectations about the future, he plans current and future consumption expenditure at the beginning of each period. Government spending on goods, services, and welfare schemes directly affects consumer behavior, since the consumer derives utility from public goods, services, and transfers in kind. For example, public roads, parks, schools, libraries, and hospitals satisfy some of the consumer's wants. We assume that although he chooses the level of consumption by participating in free markets, he takes government expenditure as given. We also assume that he takes leisure as given.[3] We start our analysis with the one-period utility function

$$u(c,l,g) = (1/\gamma)(c^{\alpha} l^{\beta} g^{1-\alpha-\beta})^{\gamma}, \tag{1}$$

[2] In particular, the share of central government expenditure in GDP in Greece grew from about 23 percent in 1980 to about 45 percent in 1990, whereas the share of private saving in GDP fell from about 22 percent during 1980-1981 to about 12.5 percent during 1982-1987 and rose again to about 20 percent during 1988-1990.

[3] Hatzinikolaou (1993) treats leisure as a choice variable, but the data indicate that the intertemporal efficiency condition (or Euler equation) for leisure is misspecified.

where c is real consumption expenditure per worker, l is leisure time per worker, g is real government expenditure per worker, and α, β, and γ are parameters. This utility function is a simple representation of preferences that are nonseparable in consumption, leisure, and government expenditure, implying that the consumer considers a geometric weighted average of these variables. If this weighted average is thought of as a composite commodity, then (1) is a constant relative risk aversion utility function, provided that $\gamma < 1$. Moreover, this specification leads to an econometrically tractable estimating equation that involves the *growth rates* of c, l, and g, which are likely to be stationary. Note that if $\alpha > 0$, $\beta > 0$, and $1 - \alpha - \beta > 0$, then this utility function is strictly increasing in c, l, and g (nonsatiation), and if $\alpha > 0$ and $1 - \alpha\gamma > 0$, then it is strictly concave in the choice variable c. Note also that if $\gamma > 0$ ($\gamma < 0$), in addition to the nonsatiation restrictions, then consumption, leisure, and government expenditure are pairwise Edgeworth complements (substitutes).[4] Clearly, if $1 - \alpha - \beta = 0$, then consumer's behavior will not directly depend on government expenditure. Assuming that an increase in g, other things equal, cannot decrease utility, the statistical hypothesis we aim to test is H_0: $1 - \alpha - \beta = 0$ against H_1: $1 - \alpha - \beta > 0$.

In each period t, the consumer maximizes his lifetime utility function

$$U = E_t \sum_{s=0}^{N} [(1+\delta)^{-s} (1/\gamma)(c_{t+s}^{\alpha} l_{t+s}^{\beta} g_{t+s}^{1-\alpha-\beta})^{\gamma}] \qquad (2)$$

subject to his lifetime budget constraint

$$a_t + [(w_t(T - l_t) - c_t)] + E_t \sum_{s=1}^{N} \left\{ [w_{t+s}(T - l_{t+s}) - c_{t+s}] / \prod_{j=0}^{s-1} (1 + r_{t+j}) \right\} = 0, \quad (3)$$

[4] Two goods are Edgeworth complements (substitutes) if the marginal utility of one increases (decreases) as the quantity of the other increases.

where E_t is the mathematical expectation conditional on information available at time t; N is the length of the consumer's planning horizon; δ is a subjective discount rate, assumed to be a positive constant; a_t is the consumer's real financial wealth (measured in consumption units) at the beginning of period t; T is the time endowment (in hours); w_t is the hourly real wage (in consumption units); r_t is a real interest rate; and $T - l_t = L_t$ is work effort in period t. We assume time-separable preferences because we use annual data to estimate the model, so substitutability between goods in different periods is limited [see, e.g., Mankiw, *et al*. (1985, p. 237)]. Equation (3) says that the sum of current wealth and current and (discounted) future dissaving is zero, implying that the consumer leaves no bequests. Using Hall's (1978, p. 986) perturbation argument, the first-order condition is given by[5]

$$(1+\delta)^{-1} E_t[(1+r_t)(c_{t+1}/c_t)^{\alpha\gamma-1}(l_{t+1}/l_t)^{\beta\gamma}(g_{t+1}/g_t)^{(1-\alpha-\beta)\gamma}] = 1. \tag{4}$$

To interpret equation (4), we first calculate from equation (2) the marginal rate of substitution

$$MRS_{c_t}^{c_{t+1}} = (\partial U/\partial c_{t+1})/(\partial U/\partial c_t) = (1+\delta)^{-1} E_t[(c_{t+1}/c_t)^{\alpha\gamma-1}(l_{t+1}/l_t)^{\beta\gamma}(g_{t+1}/g_t)^{(1-\alpha-\beta)\gamma}]. \tag{5}$$

Equation (4) is the intertemporal efficiency condition, or Euler equation, for consumption. It says that when the consumer is at the optimum, he cannot become better off by transferring one unit of consumption from time t to time t+1 (by saving) or from time t+1 to time t (by dissaving). It requires that at the optimum the conditional expectation of the marginal benefit from the

[5] Equation (5) is derived in an appendix that is available from the authors upon request. Note also that as long as the one-period utility function (1) is strictly concave in c and $[1/(1+\delta)]^s > 0$, for all s, strict concavity of (1) in c implies strict concavity of (2) in current and future values of c. It then follows that the first-order condition (4) determines a *maximum*.

consumption of an additional unit in period t+1 be equal to its marginal cost. The former is measured by $MRS_{c_t}^{c_{t+1}}$ [given by equation (5)], whereas the latter is measured by $1/(1+r_t)$.[6] Equation (4) can be expected to hold regardless of labor market constraints, as long as quantity constraints are either not binding or absent from the goods and capital markets, so that intertemporal substitution in consumption is feasible.

2.2 Behavior Toward Risk

We measure risk aversion by the insurance or risk premium, φ, which is the amount of private consumption that the consumer would give up to avoid a fair gamble. If he is a risk averter, then φ is positive. To derive an expression for φ, let ε, ζ, and η be the nonsystematic components of c, l, and g, respectively, and assume that $E(\varepsilon) = E(\zeta) = E(\eta) = 0$. Assume also that ε, ζ, and η take on small values, so that moments higher than their variances, denoted by $\sigma_\varepsilon^2, \sigma_\zeta^2$, and σ_η^2, approach zero. Finally, assume that the consumer insures only against random deviations from the systematic component of private consumption. Now following Pratt (1964), let φ be the solution to $u(c - \varphi, l, g) = E[u(c + \varepsilon, l + \zeta, g + \eta)]$. An approximate solution can be obtained by using a Taylor approximation on each side of this equation. Since ε, ζ, and η take on small values (smaller than 1 in absolute value), φ must take on even smaller values, so we can approximate the left-hand side of this equation by a first-degree Taylor expansion and the right-hand side by a second-degree Taylor expansion about the current bundle (c, l, g). Setting the two approximations equal, as required by the above equation, and then solving for φ yields

[6] Suppose, for example, that $r_t = 1$, so if the consumer gives up half a unit of current consumption and increases saving by that amount, then he will be able to consume one extra unit in period t+1. The marginal cost, which is half a unit of current consumption, can be measured by $1/(1 + r_t) = 0.5$.

$$\varphi = -[1/(2u_c)](u_{cc}\sigma_\varepsilon^2 + u_{ll}\sigma_\zeta^2 + u_{gg}\sigma_\eta^2 + 2u_{cl}\sigma_{\varepsilon\zeta} + 2u_{cg}\sigma_{\varepsilon\eta} + 2u_{lg}\sigma_{\zeta\eta}), \qquad (6)$$

where subscripts of u denote first- and second-order partial derivatives of the utility function (1), and σ_{ij} denotes the covariance between the random variables i and j. Equation (6) is a special case of Paroush's (1975) equation (4), where there are as many risk premiums as commodities. It should also be noted that if l and g do not enter the utility function, then equation (6) reduces to Pratt's (1964) equation (5). For the utility function (1), equation (6) gives a specific expression, whose derivative with respect to g is given by

$$\partial\varphi / \partial g_t = [(1-\alpha-\beta)\sigma_\eta^2 / g_t^2]\{[(1-\alpha-\beta)\gamma - 1]c_t/(\alpha g_t) + (\gamma\sigma_\varepsilon\rho_{\varepsilon\eta}/\sigma_\eta) + (\beta\gamma\sigma_\zeta\rho_{\zeta\eta}c_t)/(\alpha\sigma_\eta l_t)\}, \qquad (7)$$

where ρ_{ij} is the correlation coefficient between i and j. Under H_0: $1 - \alpha - \beta = 0$ we get $\partial\varphi / \partial g_t = 0$, but under H_1: $1 - \alpha - \beta > 0$, the sign of the term in the braces in equation (7) determines the sign of $\partial\varphi / \partial g_t$, which is an empirical question and will be examined in section 3. Note that equation (7) shows only the effect of a change in the *current* level of g on φ and cannot be used to measure the effect of an announced and credible *permanent* change in government expenditure on risk aversion, since it does not involve future levels of g.

2.3 Time Preference

Fisher (1930, pp. 62, 510) defines time preference, or "human impatience," as "the (percentage) excess of the present marginal want for one more unit of *present* goods over the *present* marginal want for one more unit of *future* goods." (Fisher's emphasis.) Using this definition, time preference for consumption (*TPC*) can be approximated by the logarithm of the marginal rate of substitution of

c_t for c_{t+1} ($MRS^{c_t}_{c_{t+1}}$) evaluated at $c_t = c_{t+1}$ [see Obstfeld (1990, p. 51)].[7] Since it is not necessary to assume uncertainty when measuring time preference, we can ignore the expectations operator (E_t) and measure $MRS^{c_t}_{c_{t+1}}$ by the reciprocal of the right-hand side of equation (5). Thus,

$$TPC \approx \log(MRS^{c_t}_{c_{t+1}})\big|_{c_t=c_{t+1}} = \log(1+\delta) + \beta\gamma[\log(l_t) - \log(l_{t+1})] + (1-\alpha-\beta)\gamma[\log(g_t) - \log(g_{t+1})]. \quad (8)$$

From equation (8) it is clear that under H_0: $1 - \alpha - \beta = 0$, TPC does not depend on g, but under H_1: $1 - \alpha - \beta > 0$, an increase in g_t, holding g_{t+1} at its trend value, increases (decreases) preference for present relative to future consumption if $\gamma > 0$ ($\gamma < 0$). If the government announces that the increase in g_t will be accompanied by a proportional increase in g_{t+1}, however, then the last term in equation (8) will vanish, so there will be no effect on time preference.

2.4 Intertemporal Substitution

We measure the effect of government expenditure on the consumer's response to a change in r_t by the elasticity of intertemporal substitution in consumption (ψ_c), which is defined as the percentage change in c_{t+1}/c_t over the percentage change in $1+r_t$ [see Mankiw, *et al*. (1985), p. 232]. Ignoring E_t again, equations (4) and (5) imply that at the optimum $1+r_t = MRS^{c_t}_{c_{t+1}}$. Thus, letting a tilde (˜) denote a percentage change, we have that

$$\psi_c = (\tilde{c}_{t+1} - \tilde{c}_t)/[(\alpha\gamma - 1)(\tilde{c}_t - \tilde{c}_{t+1}) + \beta\gamma(\tilde{l}_t - \tilde{l}_{t+1}) + (1-\alpha-\beta)\gamma(\tilde{g}_t - \tilde{g}_{t+1})]. \quad (9)$$

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[7] Let MU_{c_t} denote the marginal utility of consumption in period s. Then Fisher's definition implies that $TPC = (MU_{c_t} - MU_{c_{t+1}})/MU_{c_{t+1}} = MRS^{c_t}_{c_{t+1}} - 1$, or $MRS^{c_t}_{c_{t+1}} = 1 + TPC$. Thus, if TPC is small, then this approximation is accurate.

From equation (9) it is clear that under H_0: $1 - \alpha - \beta = 0$, ψ_c does not depend on the rate of growth of g, but under H_1: $1 - \alpha - \beta > 0$, the higher the current rate of growth of government expenditure ($\tilde{g}_t$) relative to its future rate of growth ($\tilde{g}_{t+1}$), the smaller (larger) the elasticity ψ_c will be if $\gamma > 0$ ($\gamma < 0$). For example, under H_1 and under Edgeworth complementarity ($\gamma > 0$), the consumer will respond less to changes in real interest rates, the larger the difference $\tilde{g}_t - \tilde{g}_{t+1}$, implying a negative correlation between this difference and market efficiency.

3. Econometric Analysis

Our empirical model is derived from equation (4). The transition from the theoretical to the empirical model involves removing E_t from the left-hand side and adding the (expectational) error term e_{t+1} to the right-hand side of equation (4). Let $C_{t+1} = c_{t+1}/c_t$, $H_{t+1} = l_{t+1}/l_t$, $G_{t+1} = g_{t+1}/g_t$, and $W_{t+1} = w_{t+1}/w_t$. Then C_{t+1} equals one plus the rate of growth of real consumption per worker, H_{t+1} equals one plus the rate of growth of leisure per worker, etc. Thus, the estimating equation is

$$(1+\delta)^{-1}(1+r_t)C_{t+1}^{\alpha\gamma-1}H_{t+1}^{\beta\gamma}G_{t+1}^{(1-\alpha-\beta)\gamma} - 1 = e_{t+1}. \qquad (10)$$

Under the joint hypothesis of rational expectations and of a correctly specified model, the error term e_{t+1} has zero mean, is serially uncorrelated, and uncorrelated with *any* information known at the beginning of time t, when expectations are formed. An important implication is that relevant variables dated t-h, $h = 1, 2, ...$, can be used as instruments. Note that since each observation in the sample is thought to have occurred in period t+1, the subscripts t-1, t-2, and t-3 denote two, three, and four lags, respectively, so they define permissible instruments [see Hall (1988, p. 348)]. Letting $\mathbf{z}$ be an $M\times1$ vector of instruments, the joint hypothesis implies a set of M orthogonality conditions given by $E(z_{t-h}e_{t+1}) = \mathbf{0}$. Estimation of the parameters and a test of the joint hypothesis will be based on the sample counterpart of this expectation, i.e., on the sample moments $(1/n)\Sigma_{t=h+1}^{n} z_{t-h} e_{t+1}$, where n is the number of usable observations. An efficient method that exploits the information contained in these sample moments is Hansen's (1982) generalized method of moments (GMM), provided that the number of (linearly independent) instruments (M) is at least as large as the number of parameters to be estimated (K). Whenever $M > K$ (overidentification), Hansen (1982) provides us with a statistic that can be used to test the joint hypothesis. This statistic is asymptotically distributed as chi-squared with M-K degrees of freedom and is denoted here by $J(M\text{-}K)$. It is the well-known test of overidentifying restrictions. We estimate the parameters by GMM and test the overidentifying restrictions using the Regression Analysis of Time Series (RATS) econometric software.

We use Greek annual aggregate data, 1960-1990.[8] The empirical definitions of the variables and the sources of the data are given in the appendix. Here we note that we deflate current and future values of nominal variables by *expected* prices and inflation, using a series for expected inflation that passes an unbiasedness test (see the appendix). Since the consumer makes his decisions at the *beginning* of the current period, it is evident that the nominal wage, consumption, and government expenditure should be deflated by *expected* prices, and the nominal interest rate should be deflated by *expected* inflation.[9] When we use lagged values of these variables as instruments, however, we treat them as known and deflate them by actual prices and inflation.

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[8] We would prefer to use quarterly data if they were available, because our estimates have only large-sample properties. A multi-country study might be worthwhile undertaking, however.

[9] To our knowledge, nominal variables are usually deflated by *actual* prices and inflation [see, e.g., Hall (1978), Mankiw (1981), Mankiw, *et al*. (1985)]. The reason seems to be lack of reliable data for expected prices and inflation. An exception is Hall (1988), who uses an expected real interest rate. When we deflate by actual prices and inflation in this study, we obtain some evidence of model misspecification.

Our instrument sets include lagged values of the real interest rate and the rates of growth of the real wage, real government expenditure per worker, real consumption per worker, and leisure per worker. In addition, they include a constant and a dummy variable (D) that takes on the value of zero for the years 1960-1973 and the value of one for the years 1974-1990, to account for events that have taken place since 1974 and may potentially have affected the consumer's lifetime consumption choice.[10] Given our small sample size, we use only two, three, and four lags of the instrumental variables to form three instrument sets, denoted by $NLAG = 2$, $NLAG = 3$, and $NLAG = 4$. Using our earlier notation, we have $h = 1$ when $NLAG = 2$; $h = 1, 2$ when $NLAG = 3$; and $h = 1, 2, 3$ when $NLAG = 4$. The number of observations used in estimation is $n = 28$, 27, and 26, respectively.

Table 1 reports the results from four regressions. The first regression is equation (10) estimated with $NLAG = 4$ ($M = 17$) and no parameter restrictions.[11] The last three regressions estimate equation (10) with $NLAG = 4$ ($M = 17$), $NLAG = 3$ ($M = 12$), $NLAG = 2$ ($M = 7$), respectively, and with the restriction $\gamma = 1$ imposed. For each regression, we report the following. First, the number of instruments (M) used in estimation. Second, the estimates of the parameters (t-ratios in

[10] These events include the collapse of the military regime (July 1974); the Turkish invasion of Cyprus (July 1974), which raised expectations of warfare between Greece and Turkey; the oil shocks of the 1970s; the abandonment of the drachma-U.S. dollar fixed exchange rate regime (1975); the significant reduction in the net emigration from Greece after 1974; and the socialists' victories in the elections of 1981 and 1985, which may be interpreted as a change in attitudes in favor of a larger size of government. The presence of the dummy variable D in the instrument set is important only for the results of our regression 4 (see Table 1). In that regression, if D is not an instrument, then the estimate of α is negative, although not significantly different from zero, and the estimate of β is twice as large as in the first three regressions and significantly different from zero. The other results are roughly the same as those from the first three regressions, however.

[11] We do not report unrestricted estimates with $NLAG = 2$ and $NLAG = 3$ because in these cases the estimate of α is insignificant and unstable, taking the value of 0.03 when $NLAG = 2$ and the value of 0.59 when $NLAG = 3$. Also, while the estimates of β and γ are significant when $NLAG = 2$, they are insignificant when $NLAG = 3$. The estimate of $1/(1+\delta)$ is stable and in both cases highly significant, however, and the diagnostic statistics do not indicate model misspecification.

parentheses). Third, the Durbin-Watson (*DW*) statistic. Fourth, the Ljung-Box statistic for the hypothesis that the first p autocorrelations of the residuals are equal to zero. This statistic is denoted by $Q(p)$ and is asymptotically distributed as χ_p^2. Here, p = 14 when *NLAG* = 2 and p = 13 when *NLAG* = 3 and *NLAG* = 4. Fifth, the statistic $J(M\text{-}K)$ for testing the overidentifying restrictions. Sixth, the value of the χ_1^2 statistic for the hypothesis H_0: $1 - \alpha - \beta = 0$. The significance level of each of these chi-squared statistics is given in parentheses.

The last three rows of Table I report estimates from equation (10) when the restriction $\gamma = 1$ is imposed. [Thus, the utility function (1) reduces to a single Cobb-Douglas utility function.] Our motivation for imposing this restriction is the gain in the efficiency of our estimates, which can be seen by comparing the results of row 1 (unrestricted regression) with those of row 2 (restricted regression, estimated with the same instruments as the unrestricted regression). When the restriction is imposed, the estimates of the parameters and the diagnostic statistics remain virtually the same, but the parameters α and β are estimated more precisely. The data support the restriction, since the t-statistic for the hypothesis H_0: $\gamma_0 = 1$ is only 0.13 and its power is reasonably high.[12]

Before discussing the results, the following comments are in order. First, given the small sample size, it is difficult to test the stationarity assumption, required by GMM. We note, however, that while the hypothesis of a unit root in the variables C, H, G, and W can be rejected, that hypothesis for the real interest rate, actual and expected, cannot be rejected at all the conventional levels of significance. Hansen (1982, p. 1050) and Hansen and Singleton (1982, footnote 7) suggest that the stationarity assumption could be relaxed, however. Second, the tests assume that the estimates are

[12] Assuming a two-sided test and a 5-percent level of significance, the power of the test is 0.452 if H_1: $\gamma_1 = 2$ is true; it is 0.954 if H_1: $\gamma_1 = 3$ is true; and it is almost 1.000 if H_1: $\gamma_1 = 4.44$ is true. (Given the estimate of α, $\hat{\alpha} = 0.225455$, the concavity condition will be violated if $\gamma = 4.44$.) If a 10-percent level of significance is chosen, then these values for the power are 0.507, 0.976, and almost 1.000. To calculate these probabilities, we use linear interpolation in Owen's (1965) tables. The estimated standard error of $\hat{\gamma}$ is $\hat{\sigma}_{\hat{\gamma}} = 0.5198397$, so the noncentrality parameter $[(\gamma_1 - \gamma_0)/\hat{\sigma}_{\hat{\gamma}}]$ for the three alternative hypotheses considered takes on the values of 1.92367, 3.84734, and 6.61742, respectively.

random draws from a normal distribution. For nonlinear models and for small samples, however, convergence to the normal distribution may be problematic. Third, for small samples, the optimal number of instruments is uncertain. In a Monte Carlo study, Tauchen (1986) investigates the small sample (e.g., $n = 50$ and $= 75$) properties of the GMM estimator and concludes that the smaller the instrument set and the shorter the lag length, the more reliable the estimates.

We now discuss the results of Table 1. First, all parameters are significantly different from zero, although the level of significance varies. Second, the signs and magnitudes of all parameter estimates are as expected, satisfying the conditions for concavity and nonsatiation. Third, the positive sign of the estimate of γ implies that consumption, leisure, and government expenditure are pairwise Edgeworth complements. Fourth, in all cases the overidentifying restrictions cannot be rejected, even at the 10-percent level. Fifth, the hypotheses that the error term has zero mean and is serially uncorrelated cannot be rejected.[13] Finally, the hypothesis H_0: $1 - \alpha - \beta = 0$ can be rejected in favor of H_1: $1 - \alpha - \beta > 0$ at levels of significance below 10 percent. Thus, there is some evidence that the growth of government can cause a typical consumer to become more eager to consume now rather than in the future and less responsive to changes in the real interest rate.

To determine whether government growth also affects the consumer's attitude toward risk, we must determine the sign of the term in the braces of equation (7). (The term outside the braces is positive for all sets of parameter estimates.) Thus, we must estimate the parameters σ_ε, σ_ζ, σ_η, $\rho_{\varepsilon\eta}$, and $\rho_{\zeta\eta}$ (defined in subsection 2.2). After some experimentation, we model the systematic or trend components of c, l, and g by a constant, one lagged value of the variable in question, and a time polynomial (of degree 3 for c and l, and of degree 2 for g). Then we regress each of these variables on its systematic component and take the residuals from these regressions to be the estimates of ε, ζ,

[13] The absolute value of the t-statistic for the hypothesis that the mean of the error term in equation (10) is zero is less than 0.08 in all cases, and thus we cannot reject this hypothesis.

and η. We then estimate $\hat{\sigma}_{\varepsilon}$ = 0.00001503 (billion of drachmas), $\hat{\sigma}_{\zeta}$ = 27.83 (hours), $\hat{\sigma}_{\eta}$ = 0.00001524 (billion of drachmas), $\hat{\rho}_{\varepsilon\eta}$ = 0.47, and $\hat{\rho}_{\zeta\eta}$ = -0.17.[14] Given these estimates, the term in the braces of equation (7) is negative for all the years in the sample and all the sets of parameter estimates reported in Table 1, varying between -13.9 and -3.4. Thus, there is some evidence that the growth of government can cause the consumer to become less risk averse.

4. Summary and Conclusions

Government growth may cause a typical consumer to become less risk averse, less willing to postpone consumption, and less responsive to changes in real interest rates. As a result, it may reduce private saving and the efficiency of the market mechanism, thus reducing economic growth. We construct a theoretical model of dynamic consumer behavior in order to consider this possibility. Assuming a specific utility function, we derive functions for the risk premium, the time preference, and the elasticity of intertemporal substitution in consumption. We then show that if an increase in government expenditure is only temporary, it may reduce the degree of risk aversion, increase preference for present relative to future consumption, and reduce the responsiveness of the consumer to changes in real interest rates. An announced and credible *permanent* increase in government expenditure, however, has no effect on time preference and intertemporal substitution, whereas its effect on risk aversion is uncertain. Using Greek annual aggregate data, 1960-1990, we find support for these assertions. Our analysis suggests that the Greek government might contribute to private saving stability and to market efficiency by reducing discretionary spending and by credibly announcing changes in expenditure that are intended to be permanent.

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[14] The significance levels for the two correlation coefficients are 0.009 and 0.36, respectively, so we can conclude that shocks to government expenditure are positively correlated with shocks to private consumption, but uncorrelated with shocks to leisure.

Appendix: Data Description

1. *CPI*: *Consumer Price Index*, using 1985 as a base year. Source: International Monetary Fund [International Financial Statistics (IFS), series 64].

2. π: Inflation, measured by the percentage change in *CPI.*

3. π^e: Expected inflation, defined as the mathematical expectation of π_t conditional on information available up to time t-1. We construct a series, which we denote by $\hat{\pi}^e$, by regressing π_t on *lagged* relevant variables, which include rates of growth of productivity [obtained from the Summers and Heston (1991) data base], prices of imports (IFS, series 75d), exchange rate (IFS, series ahx), and money supply (IFS, series 34). If n is the last year used in the sample period in a particular regression, we take the fitted value for that year to be an observation of π_n^e. In year n+1, the information set is expanded, so a *new* regression, with an extra observation and possibly a different set of regressors, is used to estimate π_{n+1}^e, and so on. This procedure is feasible because data for the variables used in the above regressions are actually available for the period 1950-1990. We use several diagnostic tests to ensure that these regressions are reliable forecasting tools. The series thus constructed passes an unbiasedness test. In particular, using a standard t-test, the hypothesis of a zero-mean expectational error $(\pi_t - \hat{\pi}_t^e)$ cannot be rejected (the value of the t-statistic is 0.64). Also, in the regression of π_t on an intercept and on $\hat{\pi}_t^e$, the t-ratio for the hypothesis that the intercept is equal to zero is 0.71, and the t-ratio for the hypothesis that the slope is equal to one is 1.84, which is not significant at the 5-percent level.

4. CPI_t^e: Expected *CPI*, measured by $CPI_{t\text{-}1}(1+\pi_t^e)$.

5. *WAGEM*: Wage rate in manufacturing (drachmas per hour). This is a weighted average compensation of male and female wage earners in all firms employing 10 or more persons. Source: International Labor Office (Yearbook of Labor Statistics, Table 17A, various issues).

6. *CBDP*: Interest rate offered by commercial banks to resident customers for 3-12 month demand,

time, and regular savings deposit accounts denominated in drachmas (IFS, series 60l; the observation for 1960 was constructed using Bank of Greece sources).

7. *r*: Real interest rate, measured by $CBDP - \pi^e$.

8. *TEMPL*: Total employment in the economy. It includes professional army, but not persons on compulsory military service. Using two observations from the census years 1961 and 1971 (National Statistical Service of Greece, various issues), observations for the years 1981-1990 (Bank of Greece, report of the governor, various issues), and a series on the rate of growth of total employment during the period 1967-1990 [OECD (1987a, 1991a, 1992)], we construct a new series here for the period 1960-1990. For the years 1960 and 1962-1965, we estimate total employment by regressing changes in *TEMPL* on emigration, using data for the years 1967-1990. We assume that the reason for the reduction in total employment between 1961 and 1971 was the massive emigration from Greece during that decade (UN, Demographic Yearbook).

9. *HRSM*: Hours of work per week in all manufacturing firms (International Labor Office, Yearbook of Labor Statistics, Table 12A, various issues).

10. *l*: Leisure per worker per year, measured by $5840 - 50 \times HRSM$ (hours). This definition assumes that the average worker works 50 weeks per year and that the time endowment is 5840 (= 365×16) hours per year.

11. *CO*: Consumption, defined as national private consumption expenditure on nondurables, services, and semidurable goods in millions of drachmas in current prices [OECD (National Accounts Statistics; and National Accounts, v. I and II, various issues)]. This series is obtained by subtracting expenditure on new durable goods from total national private consumption expenditure.

12. *c*: Real consumption per worker, measured by $CO \times 100/(CPI^e \times TEMPL)$.

13. *EXPR*: Government expenditure, i.e., all payments (except for amortization of the public debt) of the *central* government in billions of drachmas in current prices (IFS, series 82).

14. *g*: Real government expenditure per worker, measured by $EXPR \times 100)/(CPI^e \times TEMPL)$.

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Table 1. Estimation of Equation (10) by GMM

	M	$\hat{\alpha}$	$\hat{\beta}$	$\hat{\gamma}$	$1/(1+\hat{\delta})$	DW	$Q(p)$	$J(M\text{-}K)$	χ^2 - test for H_0: $1-\alpha-\beta=0$
1	17	$0.225^{\dagger}$ (1.64)	$0.402^{\dagger}$ (1.35)	$1.067^{\dagger}$ (2.05)	0.990^{*} (216)	2.15	14.93 (0.31)	14.39 (0.35)	2.78 (0.096)
2	17	$0.246^{\ddagger}$ (2.23)	0.366^{*} (5.38)	-	0.991^{*} (206)	2.15	15.04 (0.30)	14.34 (0.42)	14.77 (0.000)
3	12	$0.246^{\dagger}$ (1.44)	0.423^{*} (4.77)	-	0.998^{*} (138)	2.13	17.55 (0.18)	11.33 (0.25)	4.49 (0.034)
4	7	$0.286^{\dagger}$ (1.51)	0.406^{*} (3.87)	-	0.995^{*} (132)	2.06	16.73 (0.27)	7.62 (0.11)	3.80 (0.051)

Notes: (1) M is the number of instruments used in estimation; (2) the numbers in parentheses underneath the estimates of the parameters are t-ratios; (3) the numbers in parentheses underneath the chi-squared statistics are levels of significance; (4) †, ‡, * significant at the 10-, 5-, and 1-percent level, respectively.