

## Asymmetric Effects of Government Spending: Does the Level of Real Interest Rates Matter?

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*This paper empirically explores how fiscal policy (represented by increases in government spending) has asymmetric effects on economic activity across different levels of real interest rates. It suggests that the effect of fiscal policy depends on the level of real rates because the Ricardian effect is smaller at lower financing costs of fiscal policy. Using threshold vector autoregression models on U.S. data, the paper provides new evidence that expansionary government spending is more conducive to short-term growth when real rates are low. It also finds asymmetric effects on interest rates and inflation and threshold effects associated with substitution between financing methods. [JEL C32, C51, E62]*

Postwar U.S. data exhibit substantial fluctuations in real interest rates. Changes in real interest rates affect the cost of financing government spending and the burden of future fiscal consolidation. This implies that the effects of fiscal policy may depend on the level of real rates. In particular, fiscal policy may be less expansionary in an environment of high real rates because a number of channels (discussed below) limit the effectiveness of fiscal policy when the burden of financing is increased. This paper empirically examines how the effects of fiscal policy (represented by increases in government spending) differ across different levels of real rates.

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Why might the effectiveness of fiscal policy depend on the real rate? In standard dynamic general equilibrium models, government spending shocks have positive effects on output and investment through various channels, principally related to intertemporal substitution and wealth effects on labor supply. These models, however, examine the local effects of fiscal policy shocks approximated around a steady state in which the real rate is constant. In general, they abstract from the financing consequences of the shocks, assuming that the size of the fiscal spending multiplier does not depend on the method of financing. Even in dynamic models that do not satisfy Ricardian equivalence, the fiscal policy multipliers are normally quite close to those in the usual dynamic general equilibrium model with an infinite horizon (Barry and Devereux, 2003).

Recent research, however, has emphasized that the financing consequences of fiscal policy may be critical for the effectiveness of policy itself. In particular, a large literature on the possibility of contractionary effects of fiscal policy argues that, in times of large deficits and growing public debt, government spending can have a weak or even negative impact by affecting expectations about future taxes. Government spending and public debt could reach a level at which further spending causes a precipitous fall in consumption by triggering expectations of a fiscal crisis. A number of papers have modeled the way in which fiscal policy can have counterproductive outcomes through this mechanism: for example, one strand of the literature proposes the “expansionary fiscal contraction” hypothesis (Barry and Devereux, 1995 and 2003; Sutherland, 1997; and Perotti, 1999).

The level of real interest rates is critical for standard evaluations of government debt sustainability (see, for example, Reinhart, Rogoff, and Savastano, 2003). Ball, Elmendorf, and Mankiw (1998) suggest the emergence of a virtuous cycle in which low real rates and rapid growth reduce fiscal debt burden. If the return on government debt is sufficiently below the output growth rate for a sufficiently long period, the government can roll over the debt and accumulated interest without raising taxes because output will likely grow faster than the debt will accumulate. Conversely, if the output growth rate becomes low relative to the return on the debt, the debt-output ratio will increase, and eventually the government will be forced to raise taxes.

Since the level of real interest rates is important for the evaluation of the fiscal burden, it may also be important for the effectiveness of fiscal spending. A persistent shock to government spending affects the probability of hitting the upper limit on the debt ratio and thus the probability of a future fiscal adjustment. When real rates are low, a fiscal expansion that is financed by deficits rather than current taxes raises the stock of public debt, but this does not generate a significant risk of hitting the upper limit on the debt-output ratio. When real rates are high, however, the same fiscal expansion is more likely to push government debt toward the upper limit within the agents’ time horizon. Economic agents then perceive that fiscal consolidation will be necessary and expect higher future tax rates on wages and capital income. As a result, such a fiscal expansion can have a negative effect on aggregate consumption and investment. Therefore, the impact of government spending will be very different, depending on whether real rates are high enough for the economy to exceed the tolerable debt burden.

What leads to changes in real interest rates? This paper is mostly agnostic on this question. Shifts in real rates (see, for example, Garcia and Perron, 1996) can be associated with shifts in productivity or in time preferences. They can also be caused by structural events, such as changes in the monetary regime or deregulation of interest rates. Canzoneri and Dellas (1998) show that operating target procedures affect real rates in a stochastic general equilibrium model: interest rate targeting results in higher real rates than does monetary aggregate targeting.

Many recent studies have examined the effect of fiscal policy shocks based on government spending (for example, Blanchard and Perotti, 2002; Alesina and others, 2002; and Fatás and Mihov, 2003). Typically, changes in government spending are associated with changes in government debt rather than in the tax rate because government debt is managed to maintain a pattern of reasonably stable tax rates over time, although sometimes government debt and associated interest payments force the government to raise taxes.<sup>1</sup> Following this line of research, we focus exclusively on government spending in assessing the effect of fiscal policy.

To examine nonlinearities in the effect of government spending that arise from shifts in the cost of financing, we employ threshold regression methods (Tong, 1990; Choi, 1999; and Hansen, 1999 and 2000). Our model specification allows government spending shocks to have different effects on economic activity, depending on the level of real interest rates. The results obtained using U.S. time series data suggest that asymmetry in fiscal policy effects is associated with nonlinearity in the behavior of investment growth, output growth, and interest rates across different levels of real rates. Linearity testing supports the existence of a double threshold (that is, three regimes), and impulse-response analysis reveals pronounced asymmetries in the dynamic response of the economy to a government spending shock.

We provide new evidence that expansionary government spending is conducive to stimulating growth in the short run when an economy faces comparatively low real interest rates. We also find asymmetric effects of government spending on nominal and real rates of interest: there are positive effects on nominal interest rates when real interest rates are low, but negative effects on nominal and real interest rates when interest rates are sufficiently high, which is in accord with earlier studies (Evans, 1985 and 1987; and Mankiw, 1987). In addition, we find positive inflation responses to government spending only when real interest rates are sufficiently low, which reconciles the cross-country evidence of recent studies on the association of fiscal balance and inflation (Koelln, Rush, and Waldo, 1996; Fischer, Sahay, and Végh, 2002; and Catão and Terrones, 2003). Furthermore, we provide some evidence on threshold effects associated with substitution between government debt and money for financing government spending.

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<sup>1</sup>Davig, Leeper, and Chung (2004) suggest that tax policy reactions can shift between periods when taxes are adjusted in response to government indebtedness and periods when other priorities drive tax decisions. We focus on government spending and its implications for future tax liabilities but not on the tax policy behavior itself.

## I. Relevant Literature on the Effectiveness of Fiscal Policy

There are numerous different theoretical approaches to the analysis of fiscal policy. In a simple Keynesian framework with price rigidity, a fiscal expansion has a multiplier effect on output. If government spending increases interest rates, however, then “crowding-out” dampens the multiplier effect.<sup>2</sup> Although fiscal policy analysis traditionally focuses on its demand-side effects, there could be supply-side effects that add to the effectiveness of fiscal policy.<sup>3</sup> Government spending on investment-type goods helps augment production capacity and thus tends to increase the fiscal multiplier by ameliorating the crowding-out effect.

On the other hand, the neoclassical approach to fiscal policy suggests that the effects of government spending stem mainly from intertemporal substitution and wealth effects (Barro, 1981; Aiyagari, Christiano, and Eichenbaum, 1992; Christiano and Eichenbaum, 1992; and Baxter and King, 1993). The wealth effect through a Ricardian channel is operative as long as increases in government spending today imply increases in current or future taxes. The resulting fall in wealth reduces consumer demand, increases labor supply, and lowers interest rates (Devereux and Love, 1995; and Barry and Devereux, 1995), and the increase in labor supply in turn increases the marginal productivity of capital and spurs investment (Ramey and Shapiro, 1998; and Burnside, Eichenbaum, and Fisher, 2003). The size of the wealth effect depends on whether the change in government spending has purely transitory or persistent effects.

Ricardian equivalence (Barro, 1974; and Evans, 1988) suggests that increased government spending will create future liabilities anticipated by agents. Empirical studies, however, have provided mixed results for Ricardian equivalence (for listings of studies, see Evans, 1988; and Hemming, Kell, and Mahfouz, 2002). If the private sector does not fully account for or discount the future taxes implied by increased government spending because agents have a short time horizon, less than perfect foresight, or binding borrowing constraints, Ricardian effects are only partial, or consumer spending will be dependent on current income (for example, Blanchard, 1985; Campbell and Mankiw, 1989; and Mankiw, 2000). In this case, fiscal policy can retain a stabilization role, and the issue of its effectiveness remains. Campbell and Mankiw (1989), using time series analysis, find the quantitative importance of “rule-of-thumb consumers”—those who do not borrow or save but consume their current income fully—possibly reflecting that numerous consumers facing binding borrowing constraints cannot engage in the intertemporal consumption-smoothing. Findings from micro data support a strong influence of current income

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<sup>2</sup>In an open economy, higher interest rates induce capital inflows and real exchange rate appreciations, which result in a deteriorating current account and offset the increase in domestic demand arising from a fiscal expansion.

<sup>3</sup>Since public services can be considered as an input to private production, government spending on public goods and infrastructure can lead to faster economic growth (Aschauer, 1989; Barro and Sala-i-Martin, 1992; and Tanzi and Zee, 1997). Such supply-side effects of fiscal policy are regarded as more important over the longer term.

over consumption.<sup>4</sup> Accounting for the role of such rule-of-thumb consumers or “spenders,” Mankiw (2000) suggests that temporary tax changes have large effects on consumption. More recently, Galí, López-Salido, and Vallés (2005), using a new Keynesian model with sticky prices, show that government spending has an expansionary effect on consumption when rule-of-thumb consumers coexist with conventional infinite-horizon Ricardian consumers.

Contrary to both the conventional Keynesian and neoclassical views, some papers suggest that fiscal spending can have negative impacts on real activity. The “expansionary fiscal contraction” hypothesis suggests that fiscal contractions can, through their impact on expectations, lead to growth in consumption and investment. In this hypothesis, a large or persistent fiscal contraction, after a prevailing expansionary fiscal stance, signals the government’s adjustment that has been delayed (Barry and Devereux, 1995; Sutherland, 1997; and Perotti, 1999). Such episodic contractions are more likely to happen in the economies that need a fiscal adjustment (for the listing of related studies, see Alesina and Perotti, 1997; Giavazzi, Jappelli, and Pagano, 2000; Hemming, Kell, and Mahfouz, 2002; and Alesina and others, 2002).

These studies emphasize that the effects of fiscal policy may depend on the state of the economy. Two strands of studies emphasize nonlinearity in the effect of fiscal policy. One strand focuses on the different characteristics of fiscal impetus. Giavazzi, Jappelli, and Pagano (2000) suggest that nonlinear effects (on saving) are associated with large and persistent fiscal impetus for industrial and developing countries, whereas Alesina and others (2002) find little evidence for different impacts of government spending (on investment) during large fiscal adjustments rather than in normal times. Bayoumi and Masson (1998), using Canadian data, show that national fiscal stabilizers have different impacts than local fiscal stabilizers because of their different implications for future tax liabilities. The second strand of studies emphasizes expectations about fiscal adjustment for debt sustainability. Bertola and Drazen (1993) suggest that, as government spending approaches a critically high level, a nonlinear relationship arises between government spending and private consumption, consistent with the expansionary effect of large fiscal cuts as part of stabilization programs. Sutherland (1997) theoretically and Perotti (1999) empirically examine how the effect of fiscal policy depends on the level of public debt, extreme values of which trigger consumers’ expectations of an increase in their future tax liability.

A fiscal expansion may also have effects on interest rates and inflation. Contrary to the hypothesis that higher interest rates caused by the fiscal expansion would have a crowding-out effect, Evans (1987) finds no positive association between budget deficits and real or nominal rates of interest, consistent with Ricardian

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<sup>4</sup>Findings from micro data help explain why consumption is strongly associated with current income: consumption is affected by anticipated tax refunds (Shapiro and Slemrod, 1995) or predictable income changes resulting from Social Security taxes (Parker, 1999); and a substantial fraction of households have near-zero net worth (Wolff, 1998), implying that many consumers do not engage in the intertemporal consumption-smoothing (Mankiw, 2000).

equivalence. Mankiw (1987) argues that an increase in government spending depresses the real interest rate because it reduces private consumption (through a wealth effect) and increases the marginal utility of consumption, which lowers the marginal rate of substitution and thus the marginal productivity of capital (through capital accumulation). Government spending can alternatively be financed with money creation which may lead to inflation, especially by governments running persistent deficits. However, Dwyer (1982) finds no evidence that higher current or past budget deficits raise the price level. Recent analyses of cross-country data suggest that the positive association between fiscal deficits and inflation is strong among high-inflation and developing countries but not among low-inflation and industrial economies (Fischer, Sahay, and Végh, 2002; and Catão and Terrones, 2003).

Many empirical studies examine the effect of government spending that is not related to the current state of the economy and thus is less prone to simultaneity problems. The empirical results on the effect of such spending are rather mixed. Ramey and Shapiro (1998), using postwar U.S. data, find that a military buildup decreases consumption and increases (nonresidential) investment, a finding consistent with neoclassical models. Blanchard and Perotti (2002), using a structural vector autoregression approach with U.S. data, show that a government spending shock has a positive effect on output and consumption but a negative effect on investment. Alesina and others (2002), using panel data from industrial countries, find that spending shocks have a negative effect on investment, a finding consistent with non-Keynesian effects of fiscal adjustment. Perotti (2004) provides evidence on the decline in the potency of government spending over the last 20 years for a group of industrial countries. Fatás and Mihov (2003), using panel data from a large set of countries, find that discretionary fiscal policy induces macroeconomic instability.

In this paper, we adopt a regime switching approach to capture asymmetric effects of government spending across different regimes, considering that no single approach—whether Keynesian, neoclassical, or Ricardian equivalence—can always fit the data.

## II. Empirical Model Specification and Estimation Methodology

### Asymmetric Effects of Government Spending

Our results suggest the relevance of models in which full Ricardian equivalence fails: agents are *partially* Ricardian, or a substantial fraction of agents are the *rule-of-thumb* consumers described by Campbell and Mankiw (1989). Partial Ricardians will reduce their spending in response to increased government spending to the extent that such spending affects the expectation of future tax liabilities during the time horizon they regard as relevant. If consumers are myopic enough to include government debt as part of the stock of their private wealth, government spending financed by debt may have a wealth effect (Kormendi, 1983). For rule-of-thumb consumers who bear no cost of future taxation and barely hold government debt, however, government spending has neither a Ricardian effect nor the wealth

effect.<sup>5</sup> In the presence of price rigidities, the expansionary effect of fiscal policy may lead rule-of-thumb consumers to increase consumption as fiscal policy boosts their current income—the “spenders” effect (see Mankiw, 2000; and Galí, López-Salido, and Vallés, 2005). Spending on public goods and infrastructure can also lead to higher productivity and thus higher investment—a productivity-enhancing effect (Aschauer, 1989; and Barro and Sala-i-Martin, 1992). The boosting impacts of fiscal policy, through borrowing constraints, wealth, and productivity-enhancing effects, will be offset partly or fully—depending on the underlying regime—by the adverse impact from the Ricardian effect. In addition, changes in real rates in response to fiscal policy—depending on the underlying regime—have an intertemporal substitution effect on consumption.

Debt sustainability critically depends on whether real interest rates are sufficiently lower than the rate of output growth. As noted earlier, the impact of government spending may be very different depending on whether real interest rates are high enough for the economy to exceed the tolerable debt burden. Since shifts in real interest rates are given exogenously, agents will perceive the underlying regime as prevailing for a sufficiently long period.

Our regime switching approach allows the behavior of key variables in the system and the net effect of government spending to vary over regimes. Based on the level of real rates, we classify the underlying states into a “high-rate” regime, a “moderate-rate” regime, and a “low-rate” regime. We estimate a multiple-equation system—threshold vector autoregression (TVAR) models—comprising regime-dependent, reduced-form equations, especially for government spending, consumption (or investment), output, and interest rates.

## A TVAR Model

Assuming that all variables are endogenous and that government spending depends on other variables as well as its own past values, we consider a TVAR model with three regimes in a simple, piecewise-linear form (see Tong, 1990; and Choi, 1999) as follows:

$$\begin{aligned} Y_t &= A_1 + B_1(L)Y_{t-1} + V_{1,t} & \text{if } s_t \leq \tau_L, \\ &= A_2 + B_2(L)Y_{t-1} + V_{2,t} & \text{if } \tau_L < s_t \leq \tau_U, \\ &= A_3 + B_3(L)Y_{t-1} + V_{3,t} & \text{if } s_t > \tau_U, \end{aligned} \quad (1)$$

where  $Y_t = (Y_t^1, \dots, Y_t^k)'$  is a vector of  $k$  variables;  $L$  is the lag operator;  $V_{i,t} = (\varepsilon_{i,t}^1, \varepsilon_{i,t}^2, \dots, \varepsilon_{i,t}^k)'$  is a  $k \times 1$  vector of error terms, with  $V_{i,t} \sim N(0, \Sigma_{v_i})$  for  $i=1,2,3$ ;  $s_t$  is the switching index; and the thresholds are ordered ( $\tau_L < \tau_U$ ). Coefficients, denoted by  $A_i$  and  $B_i$ , vary across regimes. Errors are assumed to be heteroscedastic across regimes. Threshold parameters,  $\tau_L$  and  $\tau_U$ , are assumed to be fixed and should be estimated.

<sup>5</sup>We are grateful to the referee for suggesting that we look at implications of borrowing constraints and rule-of-thumb consumers for the consumption response to a government spending shock.

## Nonlinearity Testing and Number of Thresholds

If one or more individual equations involve threshold effects, such effects can feed into the responses of the whole system. Especially if a transmission variable of government spending to private spending involves nonlinearity, it can have differential impacts on aggregate demand. For this reason, our nonlinearity test is based on an individual equation rather than the whole system.

Classical tests have nonstandard distributions when the threshold parameter is unknown a priori and not identified under the null hypothesis of linearity—a nuisance parameter problem, or the so-called Davies problem (Davies, 1987). Following Hansen's (1999 and 2000) approach to control for the Davies problem, we obtain a consistent estimate of the threshold parameter(s) by minimizing the sum of squared residuals of an equation over a grid set. In the context of equation (1), there are zero, one, or two thresholds. To determine the number of thresholds (and thus of regimes), we perform the likelihood-ratio test, which is nonstandard but free of nuisance parameters using  $p$ -values constructed from a bootstrap procedure. Hansen's approach also helps obtain the confidence interval for the thresholds using the likelihood-ratio statistic for tests on thresholds.

In the single-threshold case, we set the bounded grid set for  $\tau \in [\underline{\tau}, \bar{\tau}]$  so that each regime has at least 20 percent of the whole sample. The grid set is composed of 100 grids that evenly divide the range from the 20th to the 80th percentile of the empirical distribution of the switching index. In the two-thresholds case, given the first-stage threshold obtained from the estimation of the single-threshold model, the grid set for the second threshold is composed of 50 grids, which evenly divide the range of the empirical distribution of the switching index. As suggested by Bai (1997) and Hansen (1999), we use the refinement estimator to improve the efficiency of the threshold parameters by estimating threshold parameters in three stages.<sup>6</sup>

## Specification Tests for Asymmetry Across Regimes

We test the null hypothesis that the coefficients in each equation are equal across regimes. To deal with the Davies problem, we employ Hansen's (1996) procedure to approximate the unknown asymptotic distribution by simulation for testing the hypothesis. We calculate three test statistics and use simulated realizations of the chi-squared empirical processes underlying these statistics, assuming that the error term is heteroscedastic across regimes but homoscedastic within each regime. The statistics are functionals of the collection of Wald test statistics over the grid space: the supremum (*SupW*), the average (*AveW*), and the exponential average (*ExpW*) of all Wald statistics (Davies, 1987; Granger and Teräsvirta, 1993;

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<sup>6</sup>In stage 1, we estimate a single threshold ( $\tau_1$ ). In stage 2, the first-stage threshold is taken as the upper (lower) threshold if it is above the 65th (below the 35th) percentile of the switching index. The grid set for the other threshold ( $\tau_2$ ) is composed of 50 grids on the side with the longer leg of the  $\tau_1$  estimate. If the first-stage threshold is between the 35th and the 65th percentile, the grid set for  $\tau_2$  is composed of 25 grids on each side of the  $\tau_1$  estimate. In stage 3, we take the  $\tau_2$  estimate as its refinement estimator ( $\tau_2^f$ ) and repeat stage 2 to obtain the refinement estimator of  $\tau_1$  ( $\tau_1^f$ ).



and Andrews and Ploberger, 1994). Their significance levels are calculated using simulated empirical distributions of the statistics.<sup>7</sup>

In addition, we perform specification tests taking symmetry across *given subgroups* in the equation as the null hypothesis, as in Durlauf and Johnson (1995). We do this by splitting the data into subgroups based on the assumed  $\tau$  and examining whether the coefficients are equal across the subgroups. The  $\hat{\tau}$  obtained by the grid search is taken as the assumed  $\tau$ .

### Analysis of a TVAR Model

To find the threshold values of a TVAR model, we employ a grid search, which is useful because our regime switching approach is based on perfect discrimination among regimes and the likelihood function is not differentiable in threshold parameters. For the minimization of the criterion function, the threshold parameter is assumed to be restricted to a bounded set.

The grid search for our TVAR model, in accord with Pesaran and Potter (1997), works as follows. Compared with the refinement estimator of a double threshold, this approach *jointly* estimates the two threshold parameters, which are assumed to be the same across individual equations in the system. We set grids by generating a  $g$ -length row vector of the grid for  $\tau_L$  and  $\tau_U$ , respectively, for its bounded set. The pairwise combinations in  $\tau$  form a  $g \times g$  grid. We then estimate the TVAR model by least squares for each point in the grid to find the estimate  $\hat{\tau}$  that maximizes the conditional log-likelihood and implies estimates  $\hat{\mathbf{A}}, \hat{\mathbf{B}},$  and  $\hat{\Sigma}_V$ . The estimate  $\hat{\tau}$  will be consistent, as suggested in Pesaran and Potter (1997) and Hansen (1996).<sup>8</sup> Let  $\tau = (\tau_L, \tau_U)'$  and  $I_t(i: \tau)$  be indicator functions with  $I_t(1: \tau) = I_t(s_t \leq \tau_L)$ ,  $I_t(2: \tau) = I_t(\tau_L < s_t \leq \tau_U)$ , and  $I_t(3: \tau) = I_t(s_t > \tau_U)$ . The conditional log-likelihood up to a constant term is given by

$$\begin{aligned} \ln l(\mathbf{A}, \mathbf{B}, \Sigma_V, \tau) = & -\frac{1}{2} \sum_{t=1}^T \left[ \sum_i^3 I_t(i: \tau) \ln |\Sigma_{V_i}| \right] \\ & - \frac{1}{2} \sum_{i=1}^3 \left\{ I(i: \tau) (\mathbf{Y}^i - \mathbf{A}_i - \mathbf{B}_i(\mathbf{L})\mathbf{Y}_{-1}^i) \right\}' \left[ I_{N_i} \otimes \Sigma_{V_i} \right]^{-1} \left\{ I(i: \tau) (\mathbf{Y}^i - \mathbf{A}_i - \mathbf{B}_i(\mathbf{L})\mathbf{Y}_{-1}^i) \right\}, \end{aligned} \quad (2)$$

where  $\Sigma_{V_i} = \frac{1}{N_i} (\mathbf{Y}^i - \mathbf{A}_i - \mathbf{B}_i(\mathbf{L})\mathbf{Y}_{-1}^i) (\mathbf{Y}^i - \mathbf{A}_i - \mathbf{B}_i(\mathbf{L})\mathbf{Y}_{-1}^i)'$ ;  $\mathbf{Y}^i$  is the selected sample vector for regime  $i$ ;  $N_i$  is the number of observations in regime  $i$ ; and  $I_{N_i}$  is an  $N_i \times N_i$  identity matrix.

<sup>7</sup>We generate  $J$  ( $= 1,000$ ) realizations of the Wald statistics,  $\chi_T^{2j}(\tau)$  ( $j=1,2, \dots, J$ ), under the null of symmetry for each grid and construct empirical distributions for three functionals of the collection of the statistics over grid space  $\Gamma: \text{Sup}W = \tau \in \Gamma \sup \chi_T^2(\tau), \text{Ave}W = \frac{1}{\#\Gamma} \sum_{\tau \in \Gamma} \chi_T^2(\tau), \text{Exp}W = \ln \left\{ \frac{1}{\#\Gamma} \sum_{\tau \in \Gamma} \exp(\chi_T^2(\tau)/2) \right\}$ , where  $\#\Gamma$  is the number of grid points in  $\Gamma$ .

<sup>8</sup>Hansen's (1999) procedure, by minimizing the sum of squared errors in the threshold autoregressive model, enables one to compute the confidence intervals of thresholds for a *single* equation. Hansen's procedure for computing confidence intervals, however, is not readily applicable to the thresholds that are obtained by the maximization of the conditional log-likelihood for *multiple* equations.

We construct a total of 1,600 grid sets for  $\tau = (\tau_L, \tau_U)'$ , allowing for 40 grids for each threshold: the lower threshold ranges from the 20th to the 40th percentile of the empirical distribution of  $s_t$ , and the upper from the 60th to the 80th percentile. We estimate TVARs with the same lag order in all regimes by the least squares method. Then we obtain the impulse response function of the variables of interest to an orthogonal fiscal shock obtained through the Choleski decomposition of  $\Sigma_{v_t}$ , assuming that the economy stays within its initial regime. Empirical standard error bounds for the response function are obtained using the bootstrap method (Runkle, 1987) with 1,000 replications.

### III. Empirical Results

#### Data and Variable Sets for TVARs

We use U.S. quarterly time series data over the period 1959:1–2001:4. The details of the data used in this paper are described in the appendix. The switching index in period  $t$  is the lagged ex post real interest rate defined as  $rr_{t-1} = R_{t-1} - 400 \cdot (P_t/P_{t-1} - 1)$ , where  $R_t$  and  $P_t$  are the nominal interest rate and the price level, respectively, in period  $t$ . The three-month treasury bill rate is used as the nominal interest rate. Inflation is measured on the basis of the GDP deflator.

The top panel of Figure 1 depicts the ex post real rate of return on three-month treasury bills along with, for comparison, the ex ante real rate. The quarterly expected inflation is interpolated from the semiannual inflation forecast from the Livingston Survey database at the Federal Reserve Bank of Philadelphia, and the ex ante real rate is defined as the three-month treasury bill rate minus the expected inflation. The ex post rate dipped in the 1970s as inflation rose, whereas the ex ante rate also dipped but by a smaller amount, reflecting that the survey inflation forecast was much smoother than actual inflation during that period of high and volatile inflation. The decline in inflation, together with financial deregulation, in the early 1980s may have led to a sharp, common drift in the ex ante and the ex post rates, which is not closely related to the inflation process. After the mid-1980s the real rate movement became rather steady, but the rate itself remained moderately high. The ex ante rate is smoother than the ex post rate but tends to overshoot before 1980, when inflation was high, and to somewhat undershoot afterward. Choi (2002) shows that the real rate is negatively correlated with inflation when inflation persistence, or inflation itself, is below a threshold.<sup>9</sup>

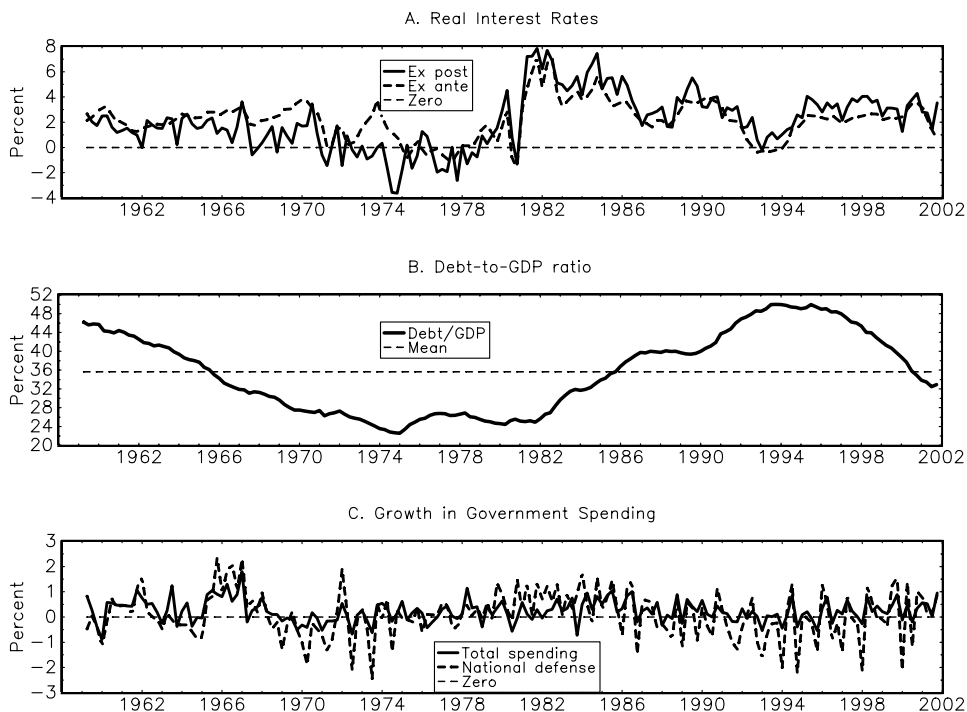
The middle panel of Figure 1 depicts the debt-output ratio, measured by nominal federal government debt divided by nominal GDP. The ratio shows a different time-varying pattern from the real rate: the debt-output ratio exhibited a downward trend in the 1960s followed by an upward trend in the 1980s and the first half of the 1990s.<sup>10</sup> The bottom panel shows growth in total government spending (real

<sup>9</sup>Inflation was rather persistent in the late 1960s and high and persistent after the 1973 oil shock. The Federal Reserve's anti-inflation policy kept inflation in check in 1982, and thereafter policy has consistently aimed at keeping inflation low.

<sup>10</sup>The regime classification according to the level of government debt (Sutherland, 1997; and Perotti, 2004) is associated with the "accumulation" of government spending that has been financed by government bond issuance, reflecting the level of "fiscal stress" of the economy.

## ASYMMETRIC EFFECTS OF GOVERNMENT SPENDING

Figure 1. Real Interest Rates, Debt-to-GDP Ratio, and Growth in Government Spending



Source: Authors' calculations.

Notes: Real rates (panel A) are measured by rates on three-month treasury bills minus expected inflation from the Livingston Survey (ex ante rate) or minus actual inflation (ex post rate). The debt-output ratio (panel B) is measured by nominal federal government debt divided by nominal GDP. The growth rates of total government spending and national defense spending (panel C) are in annualized percentages (see the appendix).

government consumption expenditures and gross investment) and in national defense spending (real national defense consumption expenditures and gross investment). Government spending grew rather rapidly in the 1960s and the mid-1980s. Growth in national defense spending is highly correlated with (correlation coefficient = 0.70), but more volatile than, growth in total spending—it is high during periods of war (especially the Vietnam War in the late 1960s) and moderately high in the 1980s (the Carter-Reagan defense buildup); it is often negative in the 1970s and the 1990s, contributing to a downward trend in the ratio of national defense spending to total government spending.

We now construct the variable sets for TVARs. In model 1, the vector  $Y_t$  includes three variables to capture real economic activity—the growth rate of real government spending ( $\Delta \ln G_t$ ), the growth rate of real private spending ( $\Delta \ln Z_t$ ), and the growth rate of real GDP ( $\Delta \ln X_t$ )—and two variables associated with financing methods and costs—the growth rate of real government debt ( $\Delta \ln D_t$ ) and the change in the (nominal) interest rate ( $\Delta R_t$ ). Real private spending ( $Z_t$ ) is measured

by either real private consumption ( $C_t$ ) or real private investment ( $I_t$ ). The interest rate is measured by the three-month treasury bill rate. Thus, model 1 is given as:<sup>11</sup>

Model 1: TVAR with the ordering  $\{\Delta \ln G_t, \Delta \ln D_t, \Delta \ln Z_t, \Delta \ln X_t, \Delta R_t\}$ .

To account for money financing of government spending and for inflation, we also consider a model with monetary growth, inflation, and the interest rate as follows:

Model 2: TVAR with the ordering  $\{\Delta \ln G_t, \Delta \ln M_t, \Delta \ln Z_t, \Delta \ln X_t, \Delta \ln P_t, R_t\}$ .

where  $\Delta \ln M_t$  and  $\Delta \ln P_t$  are, respectively, the growth rate of money and inflation. The money stock is measured by the monetary base, and the price level by the GDP deflator. The (annualized) growth rate is measured by multiplying the log difference of a variable by 400.

The periods of the low- and high-rate regimes may be partly associated with monetary policy. Tighter monetary policy, which constrains the money financing of fiscal policy, may lead to *increased* interest rates although not necessarily to a high level in the real rate (for example, tight monetary policy in the 1970s; see Choi, 1999). In particular, tighter monetary policy calls for an increase in the real rate under the Taylor rule during the Volcker-Greenspan era but not during other periods (Clarida, Galí, and Gertler, 2000). To control for a channel through which the effectiveness of fiscal policy could be affected by monetary policy, we include changes in interest rates (model 1) and interest rates and money growth (model 2). In model 1, we use all variables in first difference because we find that all level variables for the whole sample are nonstationary.<sup>12</sup> In model 2, to account for the possibility that the interest rate can be cointegrated with inflation, we include the interest rate and inflation. We set the lag length at 2 for model 1 with investment growth and at 3 for other models: this lag length selection is based on the Akaike information criterion for the whole sample VAR.

The Choleski ordering that places government spending first is based on the identifying assumption that fiscal shocks have contemporaneous impacts on, but do not respond contemporaneously to, aggregate spending and other variables (Bernanke and Blinder, 1992; and Christiano, Eichenbaum, and Evans, 1996). This assumption is likely to be a reasonable approximation because government spending must be discussed and approved before it is implemented (Alesina and

<sup>11</sup>The law of motion of government debt can be written in a simple form as  $D_{t+1} = AG_t^\alpha D_t^\beta$ . This can be rewritten in a log-differenced form,  $\Delta \ln D_{t+1} = \alpha \Delta \ln G_t + \beta \Delta \ln D_t$ , which can be extended to a more general form in a VAR.

<sup>12</sup>For the whole period, all variables in levels in model 1 are stationary, and Johansen's maximum-eigenvalue test and trace test reject the null hypothesis of cointegration in model 1 when private spending is measured by consumption but not when it is measured by investment. For model 2, the growth of monetary base and inflation have a unit root, while the ex post real rate is stationary. We find mild evidence of cointegration for model 2 when private spending is measured by investment. We also estimate TVARs in levels to account for possible cointegrations among level variables but the conclusion we obtain is qualitatively the same.

others, 2002) and reacts little to changes in macroeconomic conditions (Fatás and Mihov, 2003).<sup>13</sup>

## Tests for Threshold Effects

### *Linearity testing and number of regimes*

We perform linearity testing to determine whether a threshold effect exists. If a single regime is rejected, we then determine whether there are two or three regimes, using Hansen's (1999 and 2000) approach. In each equation, the errors are assumed to be homoscedastic within a regime but heteroscedastic across regimes.

Table 1 reports the results of linearity testing for the four key variables along with inflation, assuming a single threshold as the alternative hypothesis. The test results tend to indicate that the null of linearity is rejected in favor of a single threshold: the likelihood-ratio test statistic for a threshold effect,  $F_1$ , is highly significant for the interest rate and inflation equations ( $p < 0.01$  for both) and mildly significant for the investment growth equation ( $p < 0.10$ ) and the output growth equation ( $p < 0.15$ ). The results, however, suggest weak or little evidence of a threshold effect in the consumption growth equation (especially in model 2,  $p > 0.40$ ). The least squares estimate of the threshold  $\tau$  involves some degree of uncertainty, as indicated by a confidence interval (90 percent) that is not very tight and, in some cases, half-open, possibly owing to small sample size. The number of observations for the low- and high-rate regimes, for example, in model 1 with consumption growth, is 47 and 120, respectively, classified by a threshold of 0.945.

Although the results are not reported, the same testing procedure yielded no evidence of nonlinearity for the government spending growth equation ( $p > 0.30$  in most cases). In addition, we find little or weak evidence of nonlinearity for the government debt growth and money growth equations. The threshold estimate varies substantially both across models for the investment growth and output growth equations and across equations (the interest rate equation versus the others), suggesting the possibility of a double threshold. Since the  $F_1$  statistic tends to reject the null of no threshold effect for the investment growth, output growth (model 2), interest rate, and inflation equations, we proceed with a further test to discriminate between one and two thresholds for these equations.

Table 2 reports the likelihood-ratio test statistic of one versus two thresholds,  $F_2$ , and the refinement estimator (Bai, 1997; and Hansen, 1999) of the double threshold ( $\tau_L^r, \tau_U^r$ ) for the investment growth, output growth, interest rate, and inflation equations in each model. The  $F_2$  statistic rejects the null of one threshold in favor of two thresholds at the 5 percent level for the interest rate and inflation equations in both models. However, the  $F_2$  statistic is insignificant for the investment

<sup>13</sup>Alesina and others (2002) note that, in the United States, the yearly budget is discussed and approved during the second half of the preceding year and that additional small fiscal measures are sometimes decided during the year but most of the time become effective by the end of the year. Fatás and Mihov (2003) suggest that spending is less prone to simultaneity problems in determining fiscal policy effects than the budget deficit is because spending is not related to the current state of the economy whereas the budget deficit is largely affected by the cycle.

Table 1. Tests for a Single Threshold and Threshold Estimates

Equation and Model	Likelihood Ratio <sup>1</sup>		Threshold Value $\tau^2$		Number of Observations for Regimes	
	$F_1$	$p$ -value	Point Estimate	90% confidence interval	Low	High
Consumption growth equation						
Model 1 with $\Delta \ln C_t$	43.4	0.089	0.909	[0.349, 0.963]	45	122
Model 2 with $\Delta \ln C_t$	37.7	0.490	0.584	[ - , 0.999]	40	125
Investment growth equation						
Model 1 with $\Delta \ln I_t$	32.2	0.058	2.715	[2.516, 3.563]	103	65
Model 2 with $\Delta \ln I_t$	52.5	0.072	0.837	[ - , 1.144]	43	122
Output growth equation						
Model 1 with $\Delta \ln C_t$	39.4	0.142	0.945	[0.566, 1.180]	47	120
with $\Delta \ln I_t$	28.8	0.116	0.909	[ - , 1.144]	45	123
Model 2 with $\Delta \ln C_t$	52.7	0.058	1.126	[0.783, 1.180]	52	113
with $\Delta \ln I_t$	64.6	0.008	0.223	[ - , 0.891]	34	131
Interest rate equation						
Model 1 with $\Delta \ln C_t$	125.1	0.000	3.798	[3.780, - ]	133	34
with $\Delta \ln I_t$	87.1	0.000	3.798	[3.780, - ]	134	34
Model 2 with $\Delta \ln C_t$	91.0	0.000	3.509	[3.166, - ]	123	42
with $\Delta \ln I_t$	101.7	0.000	3.509	[3.318, - ]	123	42
Inflation equation						
Model 2 with $\Delta \ln C_t$	147.8	0.000	0.801	[0.783, 0.927]	42	123
with $\Delta \ln I_t$	165.8	0.000	0.801	[0.783, 0.927]	42	123

Source: Authors' calculations.

<sup>1</sup>The likelihood-ratio test statistic,  $F_1$ , is for the null hypothesis of no threshold against the alternative hypothesis of single threshold, following Hansen (1999 and 2000). The number of simulation replications for each grid was set at  $J=1,000$ . The lag length is set at 2 for model 1 with investment growth and at 3 for other models, based on the Akaike information criterion for the whole sample vector autoregressions.

<sup>2</sup>The grid set:  $\Gamma = \{100 \text{ grids evenly dividing the range from the 20th to the 80th percentile of the switching index}\}$ . The 90 percent confidence interval for  $\tau$  is computed using the likelihood-ratio statistic following Hansen (1999).

equation in both models and largely insignificant for the output equation, except that it is significant at the 5 percent level for model 2 with  $\Delta \ln C_t$ . The 90 percent confidence intervals for two thresholds are calculated based on the refinement estimator. Again, the confidence intervals of the threshold estimates are not very tight and, in several cases, half-open, possibly owing to small sample size. Since the  $F_2$  statistic indicates strong evidence of a double threshold for the interest rate and inflation equations, we hereafter consider a three-regime model.

### *Specification tests for asymmetry across regimes*

Table 3 summarizes the specification test results for individual equations in TVARs with a double threshold. Following Hansen (1996), we compute three test statis-

Table 2. Tests for a Double Threshold and Threshold Estimates

Equation and Model	Likelihood Ratio <sup>1</sup>		Threshold Value $\tau_f^2$						Number of Observations			
	$F_2$	Asymptotic $p$ -value	Lower bound ( $\tau_f^L$ )			Upper bound ( $\tau_f^U$ )			Low-rate regime	High-rate regime	Moderate-rate regime	
			Point estimate	90% confidence interval	Point estimate	90% confidence interval						
<b>Investment growth equation</b>												
Model 1 with $\Delta \ln I_t$	22.7	0.275	0.681	[ - , 1.111]	2.737	[2.582, 3.599]	41	62	65			
Model 2 with $\Delta \ln I_t$	45.2	0.299	0.359	[0.345, 0.808]	2.710	[2.297, 3.320]	37	63	65			
<b>Output growth equation</b>												
Model 1 with $\Delta \ln C_t$	42.5	0.116	1.113	[0.902, 1.136]	2.317	[ - , 2.508]	52	39	76			
with $\Delta \ln I_t$	21.1	0.324	0.909	[0.247, 1.116]	3.798	[3.478, - ]	47	87	34			
Model 2 with $\Delta \ln C_t$	64.9	0.035	1.125	[0.340, 0.808]	3.217	[1.548, 1.594]	52	63	50			
with $\Delta \ln I_t$	50.4	0.184	0.360	[0.337, 0.886]	3.660	[3.498, - ]	37	91	37			
<b>Interest rate equation</b>												
Model 1 with $\Delta \ln C_t$	55.8	0.025	1.376	[0.776, 2.362]	3.798	[3.783, - ]	59	74	34			
with $\Delta \ln I_t$	33.5	0.051	0.896	[0.820, 1.208]	3.798	[3.776, - ]	44	90	34			
Model 2 with $\Delta \ln C_t$	145.9	0.000	0.186	[ - , 0.309]	3.813	[3.790, - ]	32	98	33			
with $\Delta \ln I_t$	129.3	0.000	0.186	[ - , 0.274]	3.813	[3.416, - ]	32	98	33			
<b>Inflation equation</b>												
Model 2 with $\Delta \ln C_t$	100.6	0.035	0.845	[0.779, 0.911]	3.486	[3.233, 3.486]	43	78	42			
with $\Delta \ln I_t$	102.4	0.025	0.186	[ - , 0.568]	3.396	[3.233, 3.486]	32	86	45			

Source: Authors' calculations.

<sup>1</sup>The likelihood-ratio test statistic,  $F_2$ , is for the null hypothesis of single threshold against the alternative hypothesis of double threshold, following Hansen (1999). The number of simulation replications for each grid was set at  $J=1,000$ . The lag length is set at 2 for model 1 with investment growth and at 3 for other models.

<sup>2</sup>The threshold parameters are based on the refinement estimator (Bai, 1997; and Hansen 1999). The grid set:  $\Gamma = \{50 \text{ grids for the refinement estimator, given the first- or the second-stage threshold estimate}\}$ .

Table 3. Specification Tests for Asymmetry in TVARs

Equation and Model	Type of Test Statistics								
	Across regimes <sup>1</sup>				Across subsamples <sup>2</sup>				
	<i>SupW</i>	<i>p</i> -value	<i>ExpW</i>	<i>p</i> -value	<i>AveW</i>	<i>p</i> -value	$\chi^2$	<i>df</i>	<i>p</i> -value
<b>Consumption growth equation</b>									
Model 1 with $\Delta \ln C_t$	61.0	0.017	51.0	0.008	27.9	0.012	52.2	32	0.013
Model 2 with $\Delta \ln C_t$	84.6	0.000	60.9	0.006	37.1	0.000	60.7	38	0.011
<b>Investment growth equation</b>									
Model 1 with $\Delta \ln I_t$	47.9	0.017	39.6	0.007	21.2	0.013	39.4	22	0.013
Model 2 with $\Delta \ln I_t$	80.3	0.001	60.6	0.007	36.2	0.001	62.4	38	0.008
<b>Output growth equation</b>									
Model 1 with $\Delta \ln C_t$	58.6	0.026	43.6	0.052	25.7	0.032	56.2	32	0.005
with $\Delta \ln I_t$	43.0	0.056	33.0	0.049	18.1	0.054	30.2	22	0.114
Model 2 with $\Delta \ln C_t$	90.9	0.000	65.0	0.002	40.5	0.000	67.7	38	0.002
with $\Delta \ln I_t$	89.4	0.000	65.2	0.001	40.1	0.000	50.3	38	0.006
<b>Interest rate equation</b>									
Model 1 with $\Delta \ln C_t$	117.4	0.000	87.9	0.000	53.7	0.000	76.8	32	0.000
with $\Delta \ln I_t$	89.8	0.000	75.3	0.000	41.0	0.000	74.3	22	0.000
Model 2 with $\Delta \ln C_t$	110.0	0.000	81.3	0.000	50.8	0.000	100.4	38	0.000
with $\Delta \ln I_t$	118.8	0.000	93.0	0.001	54.8	0.000	110.2	38	0.000
<b>Inflation equation</b>									
Model 2 with $\Delta \ln C_t$	216.7	0.000	173.5	0.000	102.4	0.000	154.1	38	0.000
with $\Delta \ln I_t$	210.5	0.000	182.7	0.001	101.6	0.000	197.4	38	0.000

Source: Authors' calculations.

<sup>1</sup>The result for testing the null hypothesis that the coefficients in the indicated equation are the same across regimes, following Hansen (1996). The number of simulation replications for each grid was set at  $J=1,000$ . The set of grid points for  $(\tau_t, \tau_{tj})$ :  $\Gamma = \{ \text{all the grids at grid search with 40 grids for each threshold} \}$  ( $\# \Gamma=1,600$ ).

<sup>2</sup>The result of the Wald tests of the null hypothesis that the coefficients of the indicated equation are constant across the indicated subsamples grouped by  $\tau_t$ , which is reported in the corresponding notes of Figures 3 and 4. The test statistics follow the chi-square distributions with degrees of freedom ( $df$ ) under the null. The  $df$  for the interest equation in a TVAR equals  $2 \times (1 + \text{number of variables} \times \text{number of lags})$ .



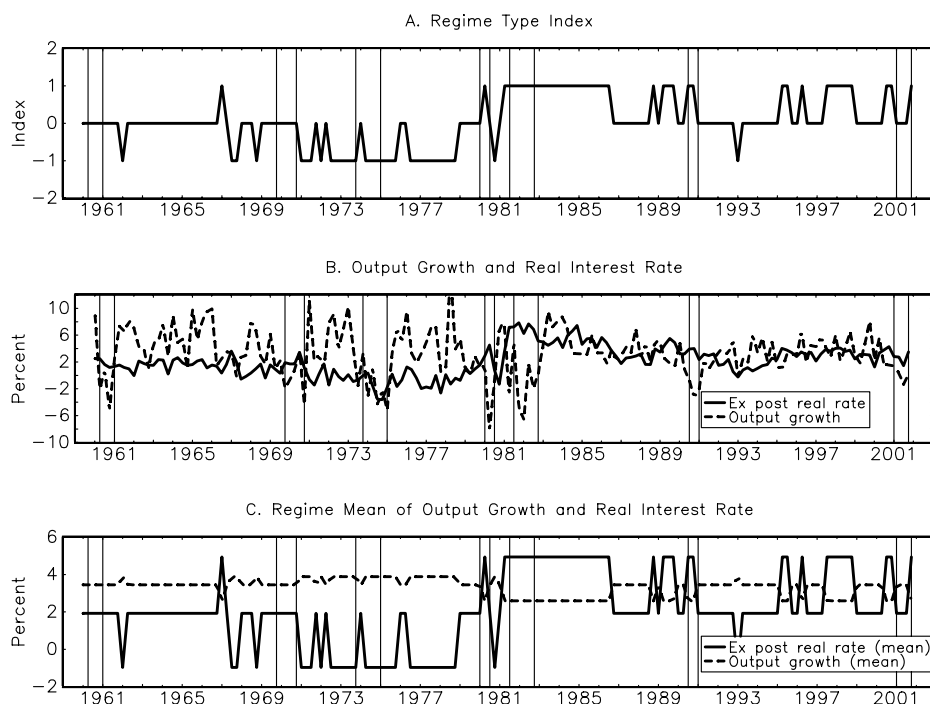
tics: *SupW*, *AveW*, and *ExpW*. All three indicate that linearity is rejected against a double threshold at the 1 percent level in all cases for the interest rate and inflation equations and at the 5 percent level or less for the consumption growth and investment growth equations. Linearity in the output growth equation is rejected for model 1 by all the statistics at the 10 percent level and for model 2 at the 1 percent level. We also performed a specification test for parameter constancy across given subsamples as if the threshold estimate  $\tau$  (reported in the notes to Figures 3 and 4) were the true value, as in Durlauf and Johnson (1995) and Choi (1999 and 2002). The last three columns of the table provide evidence against parameter constancy across the subsamples in most cases except for the output growth equations in model 1 with investment growth ( $p > 0.10$ ). Although the results are not reported, we found no evidence of asymmetry in government spending growth (test statistics were insignificant at the 10 percent level in all models). Taken together, these results give credence to threshold effects in interest rates and inflation and (sometimes mild) support for such effects in consumption growth, investment growth, and output growth.

Overall, the symmetry test results provide evidence on the coefficient shifts in most of the key regressions, consistent with the proposed switching mechanism. An increase in government spending not only directly affects aggregate demand but also affects consumption and investment through, for example, its impact on interest rates. In addition, it involves the negative impact from the Ricardian effect through anticipated increases in future tax liabilities. Weak evidence of asymmetry in consumption (as seen in Table 1) may reflect that shifts in one parameter are offset by changes in others, leaving the net effect largely intact. However, asymmetry in other equations in the system can feed into the dynamic properties of consumption through other variables in the system. Thus, and more important, we need to explore whether the dynamic responses of the economy to a government spending shock exhibit asymmetry across regimes.

Before examining the different impacts of a spending shock across regimes, we look at how the regime type and the relation between real rates and output growth evolve over time. The top panel of Figure 2 displays the regime-type index based on the estimated thresholds for model 1 with investment growth. The index indicates the low- and moderate-rate regimes until 1980 (with one exception in 1967:1) but the high- and moderate-rate regimes after 1980 (with one exception in 1993:1). The middle panel of Figure 2 depicts output growth (quarter to quarter, annualized) along with real interest rates—output growth is seen to be the more volatile series (especially before the mid-1980s). The bottom panel of Figure 2 shows that the regime mean of real interest rates is negatively related to that of output growth: the mean of the real rate is  $-0.97$  percent for the low-rate regime,  $1.92$  percent for the moderate-rate regime, and  $4.94$  percent for the high-rate regime; means for output growth are  $3.89$ ,  $3.45$ , and  $2.59$  percent, respectively.

Real interest rates are negatively associated with output growth in terms of the regime mean, and the regime-mean difference of real rates from one regime to the next is more than 2 percent, much greater than that of output growth. So abstracting from average output growth in measuring the switching index seems

Figure 2. Regime-Type Index, Real Interest Rates, and Output Growth



Source: Authors' calculations.

Notes: Regime type in panel A is determined on the basis of the estimated thresholds for model 1 with investment growth (see notes to Figure 3): low-rate regime, -1; moderate-rate regime, 0; high-rate regime, 1. The regime mean in panel C is the subsample mean of a variable for the corresponding regime. Vertical lines indicate National Bureau of Economic Research business cycle peaks and troughs.

a reasonable approximation.<sup>14</sup> In addition, business cycle recessions have no systemic relations with the classified regimes (top panel of Figure 2) although they are clearly negatively associated with output growth (middle panel). Lastly, the relationship between real rates and output growth accounts for the debt-output ratio transition in the middle panel of Figure 1: for example, the dominance of the low-rate regime before 1980—characterized by periods when, on average, real interest rates were lower than output growth rates—is associated with a downward trend in the debt-output ratio.

### TVAR Models and Impulse Responses to a Government Spending Shock

We consider a positive shock of 1 percentage point to the rate of government spending growth (annualized) and its dynamic effects on the variables in TVARs. Our

<sup>14</sup>As a result, the use of an alternative switching index—the real interest minus the regime-mean output growth—will not affect the result, while it requires an iterative estimation to obtain the regime-mean output growth.

impulse-response analysis focuses on the case in which the real rate is around its mean under each regime so that small changes in the real rate after the shock do not entail a shift to another regime.

The response of government spending growth itself to such a shock shows a similar pattern across regimes (results not shown here): the only difference, if any, is greater persistence under the low-rate regime than under the others. Therefore, we focus here on whether the responses of real activity (consumption, investment, and output) to the shock confirm empirically the anticipated larger impact of government spending when that spending creates less future liabilities than when it creates more future liabilities. We also examine the associated responses of interest rates, inflation, and financing methods.

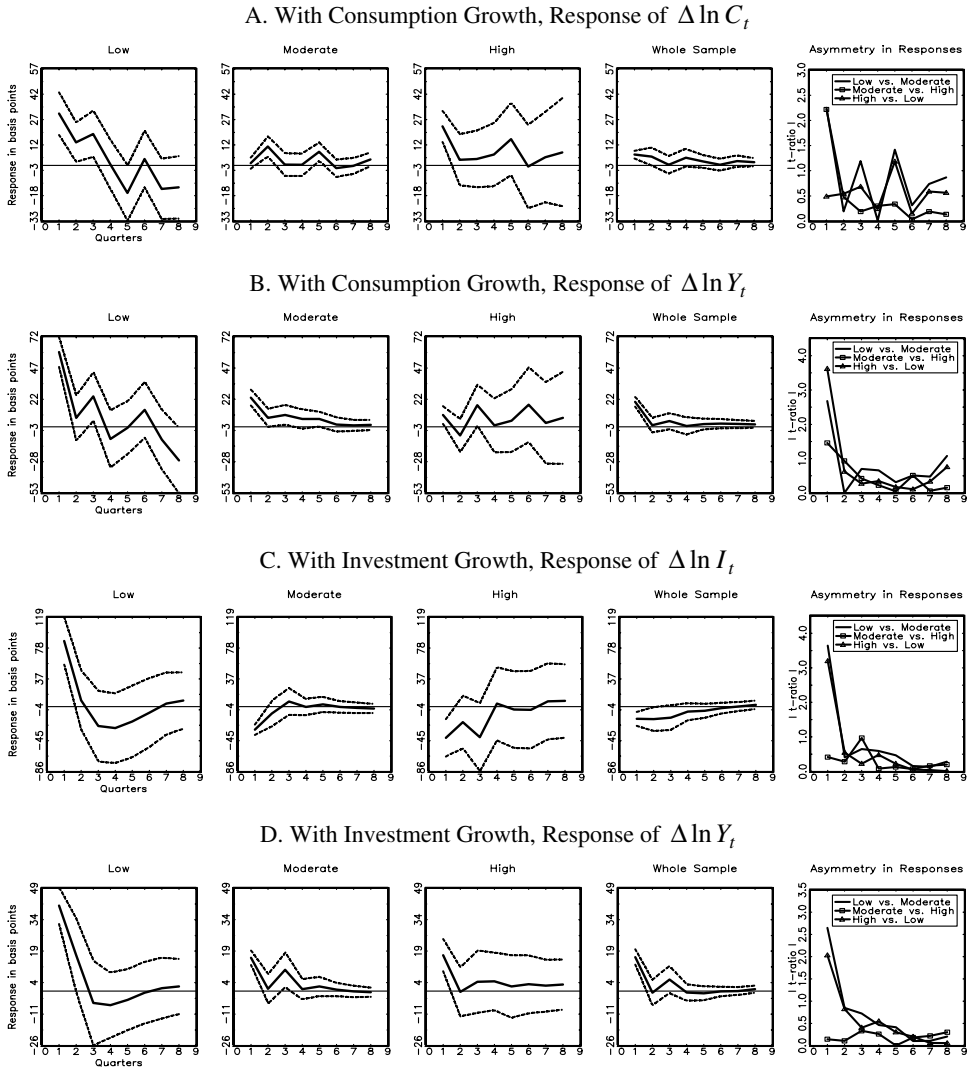
### *Responses of real activity*

Figures 3 and 4 depict the responses of consumption or investment growth and output growth in models 1 and 2. To diagnose the statistical significance of the difference in the responses between a pair of regimes, we report a  $t$ -ratio based on the estimated responses and the standard errors obtained from bootstrapping—the ratio of the difference in responses to the square root of the sum of squared simulated standard errors (regimes are independent of each other). The figures (last column) also depict the  $t$ -ratios (in absolute value) as test statistics for the null hypothesis of response equality at each horizon between two regimes for three pairs: the low-versus the moderate-rate regime, the moderate- versus the high-rate regime, and the high- versus the low-rate regime.

Consumption growth responses (top row of each figure) show initial increases in the whole sample (far right graph in each row) but differ across regimes. The initial impacts on consumption growth are stronger under the low-rate regime than under the moderate-rate regime. This finding may reflect that, under the low-rate regime, positive impacts from the wealth effect (owing to increased government debt holdings) for myopic consumers, and the spenders effect for rule-of-thumb consumers, dominate negative impacts from the Ricardian channel and the intertemporal substitution effect (with positive responses of real rates; to be shown later). Under the high-rate regime, however, consumption growth initially increases. Although in principle the Ricardian effect should be stronger in this case than under other regimes, it may be offset or dominated by positive impacts from the wealth and spenders effects (while a possible positive impact from the substitution effect would be negligible, given insignificant negative responses of real rates). The asymmetry in initial response of consumption growth is more pronounced in model 1:  $t$ -ratios greater than 2.0 at the first quarter indicate significant differences in responses between the low- and moderate-rate regimes and between the high- and low-rate regimes.

The investment growth (third row in each figure) responses of the whole sample, when significant, are negative for all models—perhaps because of crowding-out effects. In contrast, the initial impacts are significantly positive under the low-rate regime but significantly negative under other regimes. This asymmetry may be explained as follows. The productivity-enhancing effect of higher spending

Figure 3. Impulse Responses of Real Activity in Model 1

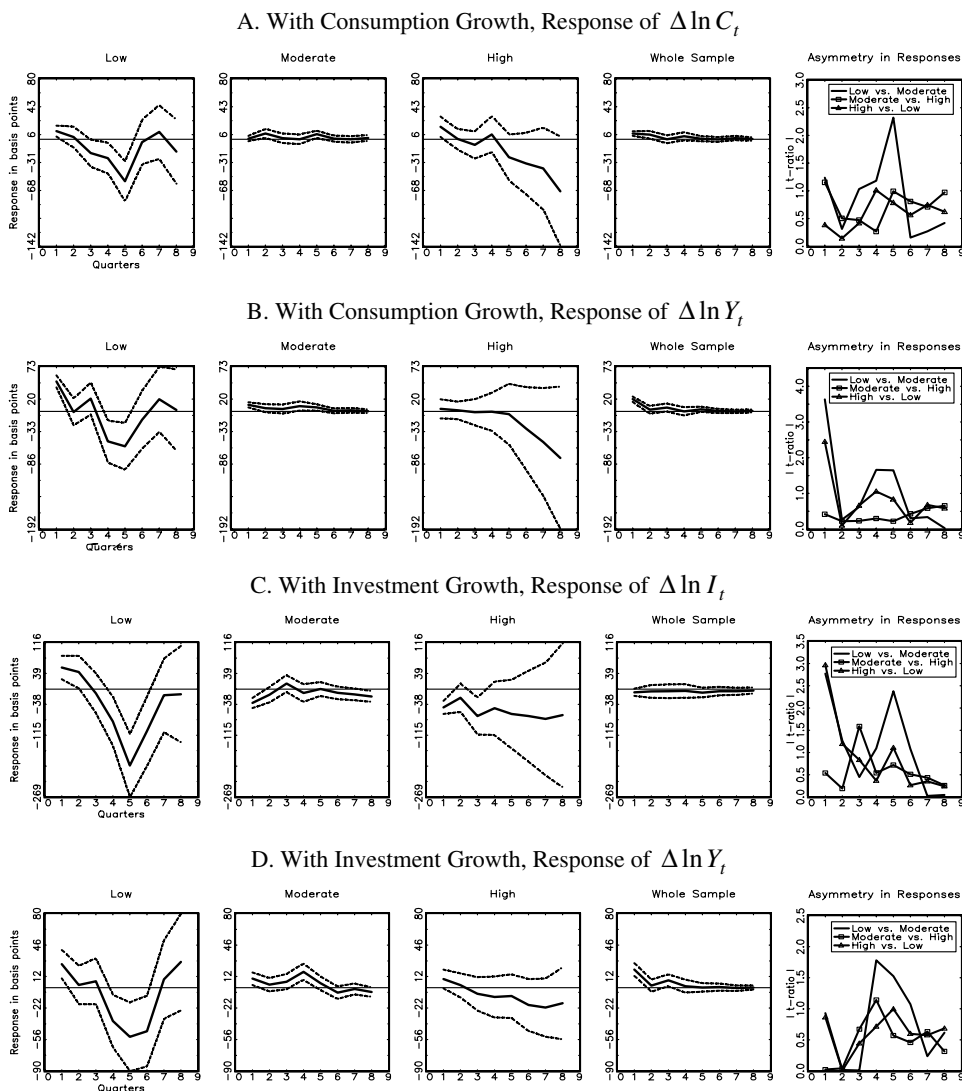


Source: Authors' calculations.

Notes: For model 1 with consumption growth, thresholds  $\tau_L$  and  $\tau_H$  are 0.358 and 3.643, respectively, and the number of observations for low, moderate, and high regimes are 37, 93, and 37, respectively. For model 1 with investment growth, thresholds  $\tau_L$  and  $\tau_H$  are 0.223 and 3.485, respectively, and the number of observations for low, moderate, and high regimes are 33, 93, and 42, respectively. Dashed lines are one-standard-error bands. The last column of each row depicts the absolute values of the  $t$ -ratio for asymmetry in responses between regimes: low versus moderate (solid line), moderate versus high (with square symbols), and high versus low (with triangle symbols).

# ASYMMETRIC EFFECTS OF GOVERNMENT SPENDING

Figure 4. Impulse Responses of Real Activity in Model 2



Source: Authors' calculations.

Notes: For model 2 with consumption growth, thresholds  $\tau_L$  and  $\tau_H$  are 0.257 and 3.705, respectively, and the number of observations for low, moderate, and high regimes are 34, 97, and 36, respectively. For model 2 with investment growth, thresholds  $\tau_L$  and  $\tau_H$  are 0.257 and 3.395, respectively, and the number of observations for low, moderate, and high regimes are 34, 88, and 45, respectively. Dashed lines are one-standard-error bands. The last column of each row depicts the absolute values of the  $t$ -ratio for asymmetry in responses between regimes: low versus moderate (solid line), moderate versus high (with square symbols), and high versus low (with triangle symbols).

induces higher investment. Under the low-rate regime, investment growth initially increases (a productivity-enhancing effect) but then declines after a few quarters as interest rates rise (a crowding-out effect). Conversely, if real rates are sufficiently high, a possible future increase in taxes has two opposing effects on investment: first, the prospect of higher corporate taxes discourages investment; second, higher income taxes decrease consumption and increase labor supply (a wealth effect) and thus investment. The negative responses of investment growth under the high-rate regime may reflect that the former effect dominates both the latter and the productivity-enhancing effect. The initial response of investment growth is significantly different between the low- and moderate-rate regimes and between the high- and low-rate regimes as shown by high  $t$ -ratios at the first quarter.

Output responses (second and fourth rows in each figure) are initially positive in most cases, but this finding is more pronounced under the low-rate regime than under other regimes. This finding is consistent with the responses of consumption and investment growth under each regime. Under the low-rate regime of model 2, output growth responses show a trough around the fifth quarter, with significantly negative values, reflecting a V-shape in the investment growth responses. The moderate-rate regime somewhat mimics the whole sample that entails an initial positive response. The initial response of output growth is significantly different between the low- and moderate-rate regimes and between the high- and low-rate regimes as shown by  $t$ -ratios greater than 2.0 at the first quarter except for model 2 with investment growth.

Table 4 reports point estimates and standard errors of averages over time of the output growth responses. Under the low-rate regime, the first-half-year response is mostly significant except for model 2 with investment growth: for example, in model 1 with consumption growth, the response is significantly positive (33.4 basis points) and about three times as large as that in the whole sample (10.6 basis points). Under the moderate-rate regime, the first-half-year response is in the range of 6 to 15 basis points but is significant only in model 1. Under the high-rate regime, the output growth response is statistically insignificant in all cases: for example, the first-half-year response is in the range of 1.2 to 8.4 basis points and insignificant.

The whole-sample analysis in Figures 3 and 4 shows initial positive responses of consumption growth and output growth and initial negative responses of investment growth, findings that are consistent with earlier empirical findings that government spending shocks have a positive effect on consumption and output (Blanchard and Perotti, 2002) and a strong negative effect on investment (Blanchard and Perotti, 2002; and Alesina and others, 2002).<sup>15</sup> In contrast, the TVAR analysis shows favorable initial effects on the growth of investment as well as the growth of consumption and output under the low-rate regime, but less favorable or (significantly or

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<sup>15</sup>Blanchard and Perotti (2002) estimate structural vector autoregression models, which contain tax, government spending, output, and an individual GDP component (such as consumption or investment) in a level form controlling for trends, for the post-1960 U.S. data. Alesina and others (2002), using a simple structural model for a panel of industrial countries, find that government spending shocks lead to a decrease in the investment-GDP ratio.

Table 4. Output Growth Responses

Quarters After Shock	Model 1				Model 2			
	Low	Moderate	High	Whole	Low	Moderate	High	Whole
	<i>With <math>\Delta \ln C_t</math></i>				<i>With <math>\Delta \ln C_t</math></i>			
1–2	33.4 (9.4)	15.2 (5.5)	1.2 (10.5)	10.6 (3.9)	23.9 (11.2)	7.9 (4.6)	2.9 (11.9)	11.3 (4.9)
3–4	7.2 (14.5)	7.8 (5.6)	9.1 (21.0)	2.6 (4.9)	-14.2 (22.7)	6.2 (5.4)	-1.1 (24.7)	3.1 (4.8)
5–8	-6.2 (10.4)	2.6 (4.4)	8.1 (33.6)	2.2 (3.0)	-12.1 (18.1)	2.3 (3.1)	-39.7 (79.7)	1.1 (2.8)
	<i>With <math>\Delta \ln I_t</math></i>				<i>With <math>\Delta \ln I_t</math></i>			
1–2	28.8 (15.5)	8.5 (3.9)	8.4 (8.5)	7.7 (3.5)	14.1 (12.3)	6.5 (4.6)	6.0 (6.6)	10.8 (4.4)
3–4	-6.2 (16.6)	5.5 (5.0)	4.5 (14.6)	2.4 (4.8)	-14.6 (20.3)	11.8 (5.8)	-8.4 (15.1)	4.8 (4.8)
5–8	-0.4 (10.8)	0.5 (2.5)	2.7 (15.3)	-0.1 (1.9)	-15.9 (17.7)	-1.6 (3.8)	-16.6 (33.0)	0.1 (2.7)

Source: Authors' calculations.

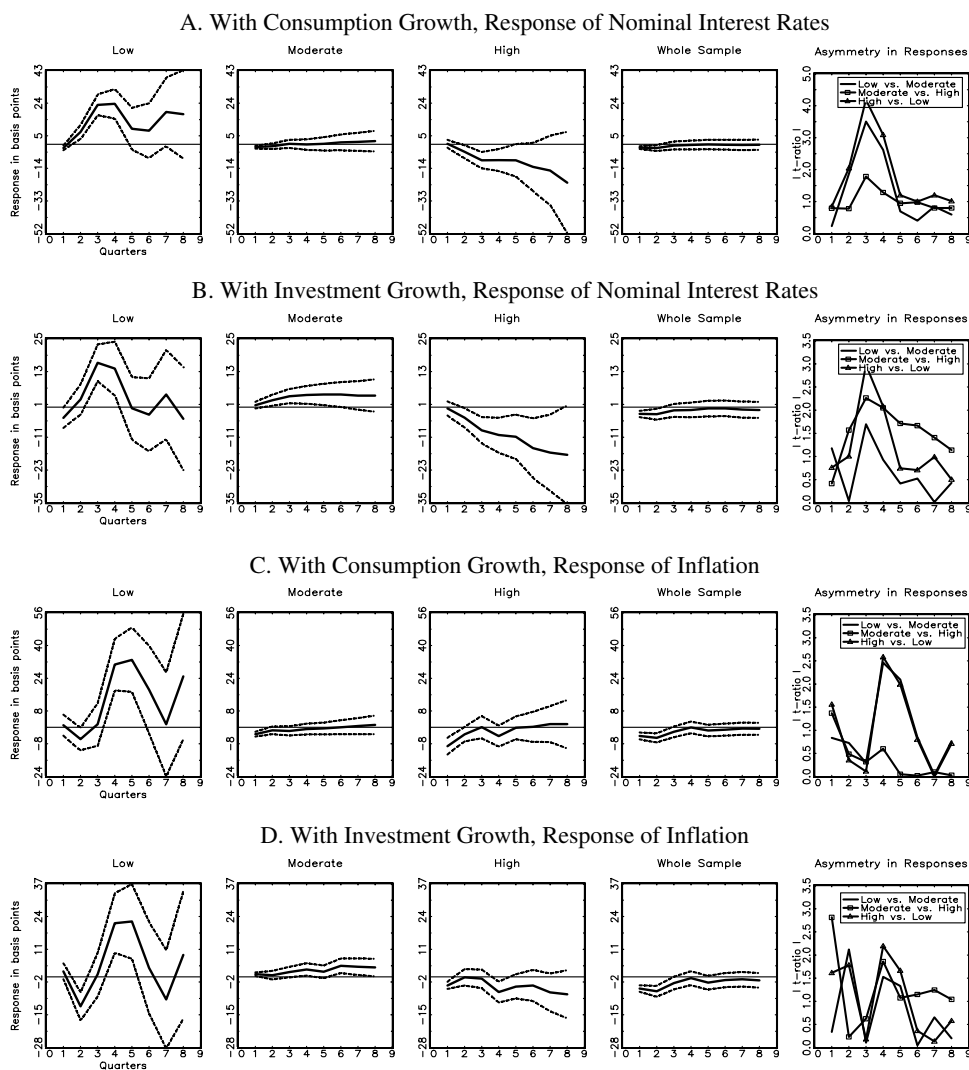
Notes: The impulse responses (in basis points) were calculated for the first half-year (1–2 quarters), the second half (3–4 quarters), and the second year (5–8 quarters) for the low-rate, moderate-rate, and high-rate regimes as well as for the whole sample. The top (bottom) panel pertains to the responses of output growth to a government spending shock in models with the consumption (investment) growth variable. Standard errors (in parentheses) are calculated from the bootstrapping with 1,000 replications. The estimated grids for TVARs are reported in the corresponding notes to Figures 3 and 4.

insignificantly) adverse initial effects under other regimes. This asymmetry indicates that expansionary government spending is more conducive to increased real activity in the short run when real rates are low than when they are high. Although no direct comparisons with existing studies are possible, this result reconciles Perotti's (2004) finding that the effects of fiscal policy on output and its components have become substantially weaker over the last 20 years for the United States (and other industrial countries) because the post-1980 period was largely one of high- and moderate-rate regimes, whereas the pre-1980 period was largely one of low- and moderate-rate regimes (bottom panel of Figure 2). Also, considering that the low-rate regime tends to be associated with the 1970s, a high-inflation era (top panel of Figure 1), the stronger effect of government spending on aggregate demand with lower real rates is in accord with Koelln, Rush, and Waldo's (1996) finding from cross-country data that the government spending multiplier is higher if inflation is sufficiently high than it is otherwise.

### *Responses of interest rates and inflation*

The top two rows of Figure 5 show the responses of nominal interest rates to the government spending shock in model 2. The nominal rate shows greater responses

Figure 5. Impulse Responses of Nominal Interest Rates and Inflation in Model 2



Source: Authors' calculations.

Notes: For model 2 with consumption growth (rows A and C), thresholds  $\tau_L$  and  $\tau_H$  are 0.257 and 3.705, respectively, and the number of observations for low, moderate, and high regimes are 34, 97, and 36, respectively. For model 2 with investment growth (rows B and D), thresholds  $\tau_L$  and  $\tau_H$  are 0.257 and 3.395, respectively, and the number of observations for low, moderate, and high regimes are 34, 88, and 45, respectively. Dashed lines are one-standard-error bands. The last column of each row depicts the absolute values of the  $t$ -ratio for asymmetry in responses between regimes: low versus moderate (solid line), moderate versus high (with square symbols), and high versus low (with triangle symbols).



under the low-rate regime than under the other regimes. Nominal rate responses are significantly positive for about three quarters from the second quarter under the low-rate regime. However, smaller positive or little responses under the moderate-rate regime and (significantly or insignificantly) negative responses under the high-rate regime and for the whole sample are shown. High  $t$ -ratios at the third and fourth quarters, especially between the high- and low-rate regimes, indicate significant differences in the nominal rate responses at the corresponding quarters.

Conventional macroeconomic theory suggests that expansionary government spending raises interest rates, a consequence that one would expect if aggregate demand rises (an income effect). But why does the shock have a positive impact on the nominal interest rate only under the low-rate regime? The initial positive responses of output growth under the low-rate regime, which exert upward pressures on interest rates, partly answer this question. Nonetheless, a thorough answer requires a further look at the responses of inflation and real interest rates.

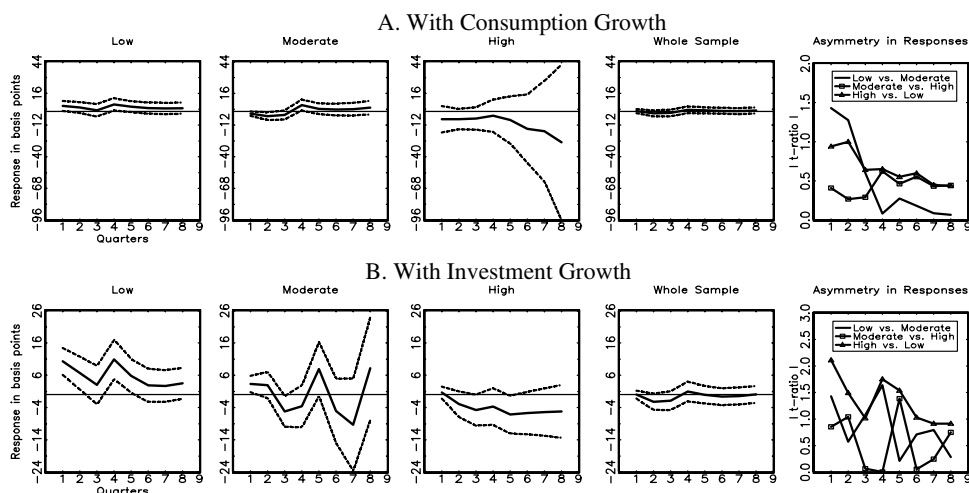
The bottom two rows of Figure 5 depict inflation responses to the shock under different regimes. Significant positive inflation responses are seen under the low-rate regime with a lag, which is largely attributable to upward pressures from aggregate demand, occurring with a lag. In contrast, under the high-rate regime and for the whole sample, negative or little responses of inflation are seen.<sup>16</sup> They are also consistent with positive (nonpositive) responses of nominal interest rates under the low-rate (high-rate) regime as implied by a Fisher effect. The asymmetry in inflation responses is significant between the high- and low-rate regimes (and between the low- and moderate-rate regimes for model 2 with consumption growth), as indicated by high  $t$ -ratios at the fourth and fifth quarters.

Nonpositive responses of nominal rates under the high- and moderate-rate regimes may suggest that the Mankiw (1987) effect of government spending on real rates is regime-dependent. The Mankiw effect will be stronger under the high-rate regime because government spending is more costly to finance and thus induces stronger negative impacts on consumption (through the Ricardian effect) and real rates when real rates are high than when they are low.

To test our conjecture that the Mankiw effect is partly responsible for the negative response of nominal rates to an increase in government spending, we look at real rate responses. We put the ex post real rate,  $rr_t$ , measured by the period- $t$  nominal rate minus the period- $t + 1$  inflation rate in place of  $R_t$  in model 2, to account for the effect of government spending on interest rates through money growth and inflation responses. We refer to this modified version as model 2'. Figure 6 shows real rate responses in this model. Under the low-rate regime, the real rate responses tend to be significantly positive, perhaps because the effect on real rates of the associated increase in aggregate demand dominates the Mankiw effect. Conversely, under the moderate- and high-rate regimes, insignificant or negative responses are shown. The whole sample shows largely insignificant effects of government

<sup>16</sup>Commodity price inflation can be included in model 2 to cope with the “price puzzle”—the finding that a monetary tightening leads to a rising rather than falling price level (Leeper, Sims, and Zha, 1996; and Christiano, Eichenbaum, and Evans, 1996). We find that the inclusion of commodity price inflation does not alter our main results.

Figure 6. Impulse Responses of Real Interest Rates in Model 2'



Source: Authors' calculations.

Notes: For model 2' with consumption growth, thresholds  $\tau_L$  and  $\tau_H$  are 1.398 and 3.812, respectively, and the number of observations for low, moderate, and high regimes are 60, 72, and 34, respectively. For model 2' with investment growth, thresholds  $\tau_L$  and  $\tau_H$  are 1.363 and 2.641, and the number of observations for low, moderate, and high regimes are 59, 42, and 65, respectively. Dashed lines are one-standard-error bands. The last column of each row depicts the absolute values of the  $t$ -ratio for asymmetry in responses between pairs of regimes: low versus moderate (solid line), moderate versus high (with square symbols), and high versus low (with triangle symbols).

spending on real rates. The real rate responses tend to be greater under the low-rate regime compared with other regimes. In particular, the initial response in the model with investment growth is significantly different between the low- and high-rate regimes, as shown by the  $t$ -ratio at the first quarter.

To the extent that a positive shock to government spending leads to budget deficits given existing tax rates, the nonpositive responses of inflation and interest rates in the whole sample and the high-rate regime are consistent with earlier findings that budget deficits have little effect on prices (Dwyer, 1982) and on real and nominal rates of interest (Evans, 1987), supported by the Ricardian effect. In contrast, the positive responses of interest rates under the low-rate regime reconcile the Keynesian prediction.

The low-rate regime tends to be associated with periods of high inflation, and the high-rate regime with periods of low inflation. Given this, our finding of positive responses of inflation only under the low-rate regime reconciles Fischer, Sahay, and Végh's (2002) finding of a significant positive association between inflation and the fiscal deficit for countries and periods with high inflation but not for low-inflation countries and for low-inflation periods in usually high-inflation countries. Also, the stronger effect of government spending on inflation with lower real rates and thus higher inflation is consistent with Ball, Mankiw, and Romer's (1988) contention that the sensitivity of inflation to aggregate demand shocks increases with inflation because economic agents adjust more frequently to keep up with inflation.

### *Responses of financing methods*

Figure 7 depicts the responses of government debt growth (in model 1) and money growth (in model 2) to the government spending shock. In all regimes and the whole sample, additional government spending initially increases debt issuance (first and second rows of the figure). Under the low-rate regime, debt finance is significantly positive initially but becomes small or insignificant as real returns on bonds, whose level was initially low, rise (see Figure 6). The responses of debt growth under the low-rate regime is different from those under other regimes—significantly in model 1 but insignificantly in model 2 (owing to large standard errors), as implied by the  $t$ -ratio at the first quarter.

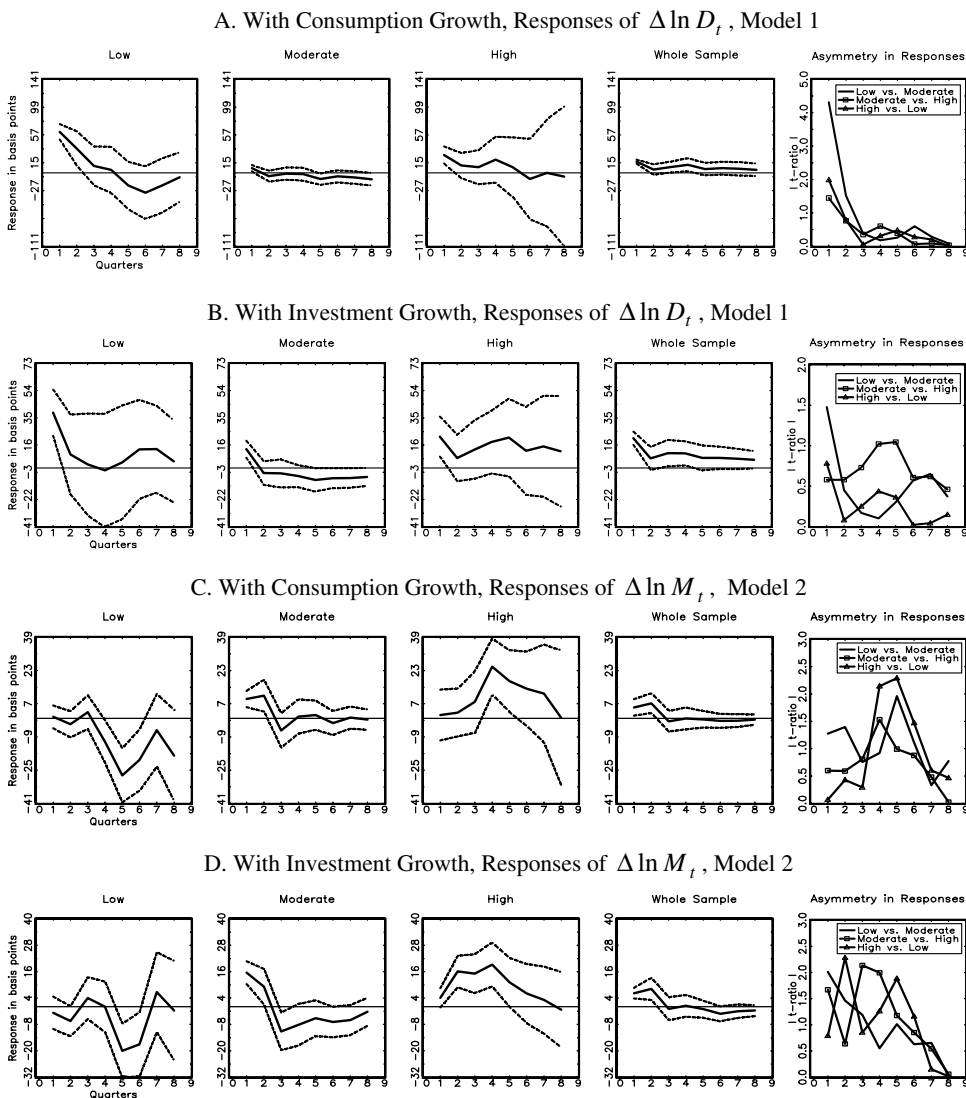
Substitution between government debt issuance and money creation in the face of different levels of real rates reflects attempts to reduce the cost of government spending. Notably, money growth (third and fourth rows) rises significantly after a short lag, showing a hump shape, under the high-rate regime but not under the low-rate regime, whereas it shows only a brief initial increase under the moderate-rate regime. This finding suggests that the financing of government spending relies on money creation only when the cost of debt financing is relatively high. Under the high-rate regime, the positive response of money growth dampens as the real rate, whose level was initially high, declines. Under the low-rate regime, the V-shaped responses of money growth seem to mirror the V-shaped responses in investment and consumption growth. The significance of asymmetry in responses is indicated by high  $t$ -ratios: for example, the  $t$ -ratios are greater than 2 between the high- and low-rate regimes at the fourth and fifth quarters in model 2 with consumption growth.

### **Robustness Checks and Discussion**

We find that the use of alternative variable sets (the real money stock M1 in place of the monetary base, and real interest rates in place of nominal interest rates) and different ordering in TVARs (for example, placing money growth after output growth) do not affect the main results qualitatively. Also, alternative lag lengths yield qualitatively similar results.

As an alternative switching index, we used the *ex ante real interest rate*. The results are similar with respect to the existence of a double threshold, but overall we find less pronounced asymmetries in the dynamic responses to a government spending shock—perhaps because the overshooting in the *ex ante* real rate before 1980 results in an obtuse discrimination of observations between the low- and moderate-rate regimes (top panel of Figure 1). Table 5 reports the point estimates and standard errors of the average output growth response over time with this switching index. A pattern similar to that in Table 4 is observed although asymmetric effects are often less pronounced. Using another alternative switching index measured by the *government debt-to-output ratio*, we found no evidence on asymmetry in consumption (or investment) and output growth equations and no evidence on asymmetric effects of government spending, although linearity testing suggests a single threshold for the interest rate and inflation equations. As noted earlier, the gradient, rather than the level, of the debt-output ratio is associated with the real interest rate. Hence, no evidence of asymmetric fiscal policy effects

Figure 7. Impulse Responses of Financing Methods in Models 1 and 2



Source: Authors' calculations.

Notes: For model 1 with consumption growth (row A) and investment growth (row B), respectively, thresholds  $\tau_L$  and  $\tau_H$  are 0.358 and 3.643, and 0.223 and 3.485, and the number of observations for low, moderate, and high regimes are 37, 93, and 37, and 33, 93, and 42. For model 2 with consumption growth (row C) and investment growth (row D), respectively, thresholds  $\tau_L$  and  $\tau_H$  are 0.257 and 3.705, and 0.257 and 3.395, and the number of observations for low, moderate, and high regimes are 34, 97, and 36, and 34, 88, and 45. Dashed lines are one-standard-error bands. The last column of each row depicts the absolute values of the  $t$ -ratio for asymmetry in responses between regimes: low versus moderate (solid line), moderate versus high (with square symbols), and high versus low (with triangle symbols).

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**Table 5. Output Growth Responses with Ex Ante Real Rates as Switching Index**

Quarters After Shock	Model 1				Model 2			
	Low	Moderate	High	Whole	Low	Moderate	High	Whole
	<i>With <math>\Delta \ln C_t</math></i>				<i>With <math>\Delta \ln C_t</math></i>			
1–2	23.2 (7.8)	–5.4 (10.4)	8.4 (13.7)	10.6 (3.9)	31.7 (12.9)	5.3 (6.3)	8.2 (6.1)	11.3 (4.9)
3–4	2.6 (11.3)	–3.3 (9.8)	14.9 (14.6)	2.6 (4.9)	–12.8 (20.2)	0.3 (6.0)	6.0 (13.1)	3.1 (4.8)
5–8	9.7 (9.0)	2.7 (5.7)	7.3 (15.9)	2.2 (3.0)	1.2 (13.5)	–1.5 (4.5)	–10.2 (32.8)	1.1 (2.8)
	<i>With <math>\Delta \ln I_t</math></i>				<i>With <math>\Delta \ln I_t</math></i>			
1–2	26.2 (9.4)	–1.1 (6.8)	1.2 (6.3)	7.7 (3.5)	31.8 (15.6)	–7.0 (5.5)	7.8 (7.2)	10.8 (4.4)
3–4	–1.1 (10.0)	–3.3 (7.2)	12.6 (11.4)	2.4 (4.8)	–14.3 (21.2)	–1.3 (7.2)	12.4 (10.0)	4.8 (4.8)
5–8	1.3 (4.3)	–1.0 (3.5)	5.6 (11.5)	–0.1 (1.9)	2.5 (17.3)	–2.7 (6.5)	4.3 (10.8)	0.1 (2.7)

Source: Authors' calculations.

Notes: The ex ante real rate is measured as the three-month treasury bill rate minus expected inflation. As a proxy for expected inflation, the six-months-ahead forecast of inflation is taken from the Livingston Survey and interpolated at quarterly frequency. See notes for Table 4. The lag lengths chosen are the same as those for models with the ex post real rate as switching index (Table 4). For model 1 with consumption growth and model 1 with investment growth, thresholds  $\tau_L$  and  $\tau_H$  are commonly 1.713 and 2.592, respectively, and the number of observations for low, moderate, and high regimes are 54, 53, and 61; and 54, 53, and 60, respectively. For model 2 with consumption growth and model 2 with investment growth, thresholds  $\tau_L$  and  $\tau_H$  are 1.157 and 2.943 and 1.157 and 2.592, respectively, and the number of observations for low, moderate, and high regimes are 37, 83, and 47; and 37, 70, and 60, respectively.

in terms of the debt-output ratio suggests that the dynamics of debt—an integral part of forming agents' expectations for future fiscal consolidation—is more important for assessing debt sustainability than the current status of debt.

The composition of expenditure may matter (Kormendi, 1983; Aschauer, 1989; Barro and Sala-i-Martin, 1992; and Tanzi and Zee, 1997). For example, an increase in spending on government wages and salaries will have a less favorable impact on output than equivalent expenditure on goods and services and capital projects. Also, a shock to defense spending may have a different impact on the economy than a shock to spending elsewhere in the budget.<sup>17</sup> However, we consider total government spending rather than expenditure composition, emphasizing the implication of the financing cost of government spending for future tax liabilities

<sup>17</sup>Kormendi (1983) finds from U.S. data that defense spending is between government investment and government consumption in terms of the size of the crowding-out effect on private consumption. Evans and Karras (1998), using cross-country data analysis, suggest that private consumption and nonmilitary government spending are substitutes, whereas private consumption and military spending are complements.

as a whole: it is difficult to take into account the implications of any individual expenditure component for the economy's tax liabilities because higher spending on any individual component could be offset by lower spending on others. Thus, using national defense spending as a measure of fiscal policy has limitations for our purpose. Nonetheless, we use this measure of fiscal policy for comparison and find that linearity testing supports threshold effects in TVARs. Compared with a shock to total government spending, a shock to national defense spending shows similar but less marked asymmetric effects across regimes and tends to have a greater crowding-out effect on consumption and investment irrespective of the regime (results not shown).

Our impulse-response analysis assumes no communication across regimes—that is, the economy stays within its initial regime—as we consider the case when the real rate is around its *mean* under each regime so that changes in the real rate after the shock do not entail a shift to another regime. This assumption remains robust as a good approximation because a 1 percentage point shock does not have much of an effect on the real rate: the effect is only about 10 basis points at its peak or trough, as Figure 6 shows. Thus, the current regime at the mean value of real rates is expected to prevail after the shock—the cumulative response of the real rate after eight quarters is less than 80 basis points, so the switching index does not hit threshold values. Nonetheless, one may consider a more general case in which a government spending shock affects the real rate enough to cause switching back and forth across regimes. For example, suppose government spending rises in the low-rate regime. Economic agents would anticipate a small rise in financing costs, and thus the crowding-out effects would be small, but in fact there is some probability of switching to a higher-rate regime.<sup>18</sup> In general, however, the (conditional and nonlinear) responses depend not only on the level of real rates (initial conditions) but also on the *size* of the shock, rendering any summary of expected responses intractable.

Since the low-rate regime is largely concentrated in the 1970s, one may wonder if a subsample period analysis may give similar implications. In general, however, time series sufficiently long for regression analysis will involve a mix of different regimes. In particular, since observations during the 1990s are rather an even mix of the high- and moderate-rate regimes (see Figure 2), threshold models can better explain the data than can the subsample analysis. Moreover, the low-rate regime, largely coinciding with a high-inflation era in the United States, corresponds to observations for countries and periods with high inflation in the cross-country dimension. As noted earlier, our finding on the positive response of inflation only under the low-rate regime reconciles Fischer, Sahay, and Végh's (2002) finding from cross-country data. Therefore, rather than restricting differential impacts of fiscal policy by using a chronological time scale, our threshold model analysis—

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<sup>18</sup>To allow for shifts to other regimes at the margin, one may consider the estimation averages over the actual histories of real rates conditional on each regime, given a *fixed* size of shocks. This approach will somewhat smooth out differences across regimes but will not affect our results qualitatively, given that the moderate-rate regime, as a middle ground in the characteristics of responses, buffers a transition from one extreme regime to the other extreme unless the shock is extremely large.

which assigns an observation to one of the regimes classified by a switching index and threshold estimates—provides implications for state- or regime-dependent effects of fiscal policy, encompassing those that otherwise might have been obtained from subsample time series or cross-section analyses.

We find evidence that, in times of low real interest rates, a fiscal expansion is conducive to boosting economic activity in the short run. However, such a stimulating effect dies out fast if fiscal policy continues to be expansionary enough to deteriorate the economic environment by raising real rates and accelerating debt accumulation, thus switching to a high-rate regime. In particular, an expansionary fiscal policy could be less than effective for an economy with persistent government deficits and pervasively high real rates. Since perfect foresight during the entire time horizon of interest is often far from reality, agents can update their assessment of debt sustainability based on the evolving status determined by the financial cost of fiscal policy. Thus, no perpetual benign effect of fiscal policy is warranted for an economy, even if it starts with the low-rate regime. Furthermore, fiscal policy entails a trade-off between volatility and efficacy because an aggressive fiscal policy induces macroeconomic volatility, which in turn lowers economic growth, as shown in Fatás and Mihov (2003).

#### IV. Conclusions

Earlier studies have looked for nonlinear effects of fiscal policy based on different characteristics of the fiscal impetus or of consumers' expectations about future fiscal adjustment to achieve government debt sustainability. Little evidence, however, has been provided on the link between the efficacy of fiscal policy and the financing cost of government spending. This paper provides new evidence on the relative effectiveness of fiscal policy at different levels of real interest rates. It shows that government spending has a significant, positive short-run impact on aggregate spending at low real rates but not much of an impact at relatively high real rates.

Additional findings on asymmetric effects of fiscal policy are noteworthy. First, government spending raises inflation and nominal interest rates, owing to higher aggregate demand, only when real interest rates are relatively low. Second, at low real rates, the effect of increased government spending on real rates is positive; and, at high real rates, it can be negative. Third, government spending induces initial debt issuances, especially at low real rates, and money creation at relatively high real rates. This indicates that substitution between debt issuance and money creation depends on the financing cost of government spending.

We interpret the new evidence on the asymmetric effects of fiscal policy on economic activity as suggesting that fiscal policy is likely to be more conducive to *short-run* growth when real interest rates are low. However, a ballooning government debt with persistent, expansionary government spending can be perceived as constraining fiscal policy—fiscal austerity thus may form a stronger foundation for the efficacy of fiscal policy in times of need.

Lastly, our findings suggest that the nonlinearity in the effects of fiscal policy is associated with the link between the time-varying cost of financing government spending and the fiscal multiplier. It will be interesting in future research to

examine how the real interest rate, which is the financing cost of government spending, is determined by, for example, the underlying monetary regime and how it therefore affects the responses of economic activity to fiscal policy in a general equilibrium framework. In doing this, as emphasized by Mankiw (2000) and Galí, López-Salido, and Vallés (2005), the substantial heterogeneity in consumer behavior, consistent with findings from micro data, may play an important role in explaining the data.

## Appendix

### Data Sources and Description of the Variables

We use the U.S. quarterly series, obtained from Federal Reserve Economic Data (FRED) at the website of the Federal Reserve Bank of St. Louis, in our analysis. Variable definitions and FRED code names are as follows:  $X$  = real GDP, chained 1996 dollars (GDPC1); nominal GDP (GDP);  $P$  = GDP deflator (= GDP/GDPC1);  $G$  = real government consumption expenditures and gross investment, chained 1996 dollars (GCEC1); real national defense spending = nominal national defense consumption expenditures and gross investment (FDEFX) divided by the GDP deflator;  $C$  = real personal consumption expenditure, chained 1996 dollars (PCECC96);  $I$  = real fixed private domestic investment, chained 1996 dollars (FPIC1);  $D$  = nominal federal government debt (defined below) divided by GDP deflator;  $M$  = the Federal Reserve Board of Governors' adjusted monetary base (BOGAMBSL); money stock M1 (M1SL); and  $R$  = the three-month treasury bill rate, percent per annum (TB3MS). The data available at monthly frequency from the source are averaged to obtain quarterly observations. The nominal federal government debt is taken from the IMF's *International Financial Statistics* and seasonally adjusted (by X12).

The growth rate of a variable  $x$  in annual percentage is defined as  $\Delta \ln x_t = 400 \cdot \ln(x_t / x_{t-1})$ . The lagged ex post real interest rate is defined as  $rr_{t-1} = R_{t-1} - 400 \cdot (P_t / P_{t-1} - 1)$ . The lagged ex ante real interest rate is measured by  $R_{t-1}$  minus the expected inflation rate for period  $t$ , for which the six-months-ahead forecast of the monthly base value of CPI taken from the Livingston Survey is interpolated at quarterly frequency.

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In statistical matter throughout this issue,

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a dash (—) indicates that the figure is zero or less than half the final digit shown, or that the item does not exist;

a single dot (.) indicates decimals;

a comma (,) separates thousands and millions;

“billion” means a thousand million; and “trillion” means a thousand billion;

a short dash (–) is used between years or months (for example, 2005–06 or January–June) to indicate a total of the years or months inclusive of the beginning and ending years or months;

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Design: Luisa Menjivar-Macdonald and Sanaa Elaroussi