A Dynamic Study of the Origins and Consequences of Changes in U.S. Corporate Taxation, 1981-1998

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We propose that U.S. corporate taxation revenue, 1981-1998, responded to political and economic forces in ways consistent with a partisan and organizational view of politics, but inconsistent with a structural dependence view. We employ Vector Error Correction (VEC) models to examine our hypotheses and offer a new method for generating impulse response function boundaries in VEC models. We find that short-term increases in corporate tax revenue followed from Democratic Presidential Administrations, decreasing business political activity, and increasing economic growth, trade, and inward foreign direct investment.

That U.S. corporate tax policy is partly determined by partisan and organizational forces as well as by domestic economic forces is a standard theoretical view. Scholars find that Democratic Presidential Administrations have differed from Republican Administrations in that the former tended to increase the tax burdens of corporate taxpayers. Other studies report that increasing business political activity has had an ameliorating effect on business tax burdens. On the economic side, corporate tax policy is counter-cyclical, investment inducing, and linked with growth and investment in long-run structural relationships.

In at least two important ways, however, the structure of the American economy differed in the 1980s and 1990s from earlier periods. First, the U.S. economy internationalized, with the oil shocks of the 1970's serving as globalization's annunciation. Second, economic agents became increasingly sophisticated in anticipating how governments would respond to economic events and how these government policies would in turn affect economic events. The rational expectations "revolution" led President Reagan, among others, to conclude that traditional U.S. macroeconomic policies were ineffective.

Some scholars argue that these structural shifts changed the dynamics of U.S. corporate tax policy. They theorize that, in open economies like the U.S., high revenue yields from corporate taxation are not sustainable. (See Gordon and Bovenberg 1996 for a review.) A more radical proposition – that U.S. corporate tax policy is unaffected by political and economic events – has come from the rational expectations framework (Williams and Collins 1997). These new propositions are allied in spirit with structural dependence theory, which holds that internationalized markets and highly sophisticated investors combine to limit the ability of democratic governments to tax capital (e.g., Dryzek 1996).

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We assess propositions regarding the origins and consequences of changes in U.S. corporate tax policy, 1981-1998. We develop theoretical and empirical reasons for believing that partisanship and organizational forces were influential in the short-term dynamics of corporate tax policy.

Our study builds on recent scholarship, and adds several new features. First, the Federal Election Commission has made available for the first time quarterly data on PAC activity, so we have established a valid quarterly time series from the first quarter of 1981 through the third quarter of 1998.[1] Second, as recommended by Williams and Collins 1997, we adopt dynamic modeling, adding partisan and international economic variables to the analyses. We use Johansen's (1995) approach to vector error correction (VEC) models because dynamic models, such as VAR models, produce misleading results unless unit roots and cointegrating relationships are accounted for, and because our time-series contained some unit roots and cointegrated relationships. Third, we present a new method for generating confidence intervals for the impulse-response functions of Vector Error Correction models. We find that U.S. corporate taxation continues to be influenced by political and economic variables in ways consistent with a partisan and organizational view of politics.

HYPOTHESES ON THE ORIGINS AND CONSEQUENCES OF CORPORATE TAXATION

Corporate Taxation and International Economic Conditions.

International trade and financial openness ought to lead to lower corporate tax revenue and downward pressure on corporate tax rates. (See, e.g., Bates and Lien 1985; Frenkel, Razin, and Sadka 1991; and Steinmo 1994.) The U.S. has become a net importer of goods, implying that U.S. firms have been less able to raise prices and profits.[2] Swank (1998) found that increasing trade openness was associated with subsequent decreases in corporate taxation.

The effects should be particularly pronounced for international financial openness. Scholars offer strong theoretical work showing that, in terms of the public interest, the optimal rate of corporate taxation in an economy open to capital flows is zero. (See Gordon and MacKie-Mason 1995 for a review.) In these open-economy models, corporate taxation has deleterious effects on growth in both investment and gross domestic product. Many economists have suggested, therefore, that the public interest is best served by eliminating corporate taxation and replacing lost revenue with other, less distorting, taxes. Further, more liberalized capital account rules worldwide encourage U.S. investors and firms either to leave for lower tax locations or to engage in "aggressive" transfer pricing. This concern is not purely academic: the Internal Revenue Code Section 482 explicitly targets "transfer pricing" that serves as a form of corporate tax evasion, and Section 6662(e)(h) penalizes misevaluation of intrafirm transfers. (Eden (1998, 383-482) summarizes the regulations.) Thus, increasing corporate taxation might deter inward, and induce outward, direct investment, thereby decreasing tax revenue.

Nonetheless, both U.S. corporate tax rates and revenue have remained high – with statutory marginal rates of 34% to 39% on taxable income over $75,000 and corporate tax collections in 1998 of $240 billion. Despite these high tax policies, Foreign Direct Investment inflows into the U.S. are truly gargantuan: $360 billion from 1981-1990 (Tang 1997, 11), and more than $1.6 trillion in 1997 alone (IMF 1999, 766). The FDI inflows into the U.S. during 1981-90 accounted for 45% of all FDI inflows among OECD countries and were more than double U.S. FDI outflows (Tang 1997, 9).

Comparative work on corporate taxation and international financial liberalization has also shown that financial openness is *not* associated with decreasing corporate tax revenue worldwide (Garrett 1995; Gordon

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[1] Robert Biersack of the Federal Election Commission provided data on quarterly registrations. (Email correspondence with Dennis Quinn, 3 June 1997, 9 June 1997, and 17 March 1999. The data are available from either the author or Dr. Biersack at biersack@fec.gov. We have quarterly Political Action Committee data from the first quarter of 1981 to the third quarter of 1998.

[2] U.S. corporate profits as a percentage of gross domestic product averaged more than 10% in the mid and late 1970s, but fell to 5% in the early 1980s. Profits in the mid-1990s averaged 8% (National Income and Product Accounts, Table 1.14, line 22).

and Bovenberg 1996). Indeed, some scholars find that increasing financial openness is associated with *increasing* corporate taxation worldwide (Quinn 1997, Swank 1998).

The apparent disconnect between theory and evidence has several explanations. One follows from the point that asset returns for some industries might be higher abroad. If U.S. firms locate only highly profitable investments within the U.S., and locate abroad investments that could be undertaken more profitably there, the increased efficiency of investments will lead to more profitable U.S. firms. U.S. corporate tax revenue will rise. A second line of explanation suggests that information asymmetries account for the "home bias" of domestic capital and for why firms do not always locate investments to the lower tax locale (Gordon and Bovenberg 1996). Finally, Gravelle (1994, 1995) demonstrates that the "zero-tax" models rely on some assumptions that are unsound; she concludes that the "imperfect substitutability of either capital or goods acts to effectively close the economy for purposes of tax incidence" (1995, 274).

A practical effect of the persistence of a high U.S. corporate tax policy is that the U.S. might serve as Stackelberg leader in enabling a worldwide system of capital taxation to persist, globalization notwithstanding (Gordon 1992). That is, the U.S. is serving as the lynchpin of a worldwide system of relatively high corporate taxation, but were the U.S. to dramatically cut its rates or revenue collections, other nations would be compelled to follow. The central role the U.S. plays makes an understanding of the determinants of U.S. tax policy, and the role of international forces in it, crucial for understanding the that persistence of high corporate tax policies elsewhere.

Investor Expectations and U.S. Corporate Taxation.

Recent work by Williams and Collins (1997) offers a fundamental critique of the scholarly empirical work on U.S. corporate taxation. They begin with the claim that "the ECTR [the effective corporate tax rate with inventory valuation adjustments and capital consumption allowances] remains relatively stable around a longer-term mean from 1977-1994…." (1997, 224). Williams and Collins hypothesize tax policy stability would follow if corporate taxation were exogenous to political and economic forces. They also hypothesized that political and economic variables respond to business taxation.

Economists often analyze corporate taxation within the context of "optimal models of tax policy" (e.g., Gordon and Bovenberg 1996, 1068-1070). Williams and Collins (1997) propose that we understand U.S. corporate tax policy in light of dynamic optimal control models, which allow for analyses of situations possibly characterized by the "time-inconsistency" problem.[3] Williams and Collins suggest that

> to minimize investment around its ideal point, Congress must set effective tax rates at their expected value. Since this value is dependent on the initial starting point, it makes sense for Congress to maintain a consistent effective tax rate. (1997, 222).

If an optimal model of tax policy accurately describes U.S. corporate tax policies, changes in the domestic economy should not matter for corporate tax policy.

Williams and Collins offer evidence that effective corporate tax rates have been unaffected by economic and political conditions and that investment is a negative response to taxation. The absence of

[3] The "time-inconsistency" problem is well explored in the economics literature regarding monetary or exchange rate policies. "Optimal" government policies might depend on the state of the economy, which changes over time, which then leads policymakers to change policies subsequently. Private economic actors learn to anticipate that governments will subsequently alter policies and adapt their behavior in anticipation of government actions. These private sector adaptations change the effects of government policies, potentially leading to the ineffectiveness of government policies.

It is not apparent that the time-inconsistency problem characterizes U.S. investment and corporate taxation policies in the same way that it might monetary or exchange rate policies (see, e.g., Gordon and Bovenberg 1996). Investment and product markets equilibrate much more slowly than do financial markets, and the ability of the U.S. Government to act quickly and subsequently enforce its will on tax policy issues on domestic firms is very different from its abilities relative to international actors in international currency markets. Furthermore, U.S. policymakers generally shape tax policy in such a way so as to lessen tax arbitrage possibilities. Changes in tax law regarding depreciation and other investment incentives generally are targeted to new investment, rather than existing investment. Tax law is frequently retroactive to the date of the law's introduction into Congress (see Ways and Means 1993, 7).

significant political economic effects on corporate taxation combined with the negative response of investment to increased tax rates lead them to the conclusion that their "model is very consistent with Lindblom's structural dependence thesis" (1997, 238).

Optimal tax models are in some sense normative, rather than descriptive. That is, given a set of assumptions about the utility functions and endowments of various actors, the optimal tax models offer guidance about what behavior *should* be optimal for investors. In a broad overview, how descriptively accurate are the propositions that tax policy is characterized by consistent tax rates and that the U.S. is structurally dependent on owners of capital?

The "Tax Green Book" (Committee on Ways and Means 1993) addressed the consistency question:

Income taxes paid by corporations now contribute about one-tenth of total revenues …[which] is a marked decline from the 20-25 percent portion they contributed during most of the 1960s. …. Some of the decline has its explanation in legislated tax reductions—increasingly generous accelerated depreciation, larger numbers of tax preferences for specific industries and assets, and reductions in the corporate tax rates. Economic trends also worked to decrease the corporate tax base; for example, many companies increasingly financed their operations with deductible debt rather than with equity. Although corporate income taxes have been raised several times since 1982, this has not been enough to reverse the long-run trend. (1993, 2.)

Case studies further suggest that the consistency proposition is questionable. Martin's (1991) study of the history of U.S. corporate tax legislation highlights both the frequency and inconsistency of major corporate tax acts. Her analysis of the intentions of the political actors sharply contrasts with the notion that Congress sets tax policy to maintain some ideal investment point through time. Boskin 1996 notes eight major revisions of the tax code with attending changes in the tax rate in the last twenty-five years, with large swings in revenue from the revisions. Tax policy may be optimal for someone, but that someone seems to change frequently. Neither tax revenue nor tax rates are characterized by consistency.

As we examine some features of U.S. tax law, we have further reasons to doubt the "exogeneity" hypothesis. The "classical" corporate tax system, which is employed by the U.S., allows for two separate accounting books, one for shareholders and one for the Internal Revenue Service (IRS). The separation of economic depreciation and the tax depreciation means that corporate tax receipts are highly likely to be affected by changes in economic events (see, e.g., Kanniainen and Södersten 1995, 428-29). Indeed, the standard view of the U.S. corporate tax system is that it has counter-cyclical and investment-inducing properties (e.g., Pechman 1987), with receipts responding positively to changes in growth and negatively to changes in investment.

As for structural dependence theory, are government officials concerned with the time-inconsistency problem because the U.S. is structurally dependent on the owners of capital? If so, the President and members of Congress have a most unusual way of showing their dependence. The U.S. has consistently "double-taxed" income from equity-base corporate investment, and this tax falls disproportionately on owners of capital (see, e.g., Gravelle 1994; Sørensen 1995). The U.S. is among the few advanced industrial nations to "double tax" the income from equity-based corporate investment.[4] Arlen and Weiss fulminate that the "two-tier taxation is unusual, unfair, and inefficient. The ill effects of the double tax are well known in Washington. Congress regularly considers legislation to eliminate the double tax …. [which] invariably die[s] a quiet death" (1995, 326-7). The continued double-taxation of corporate income is hard to reconcile with a vision of a structurally dependent United States.

The Effects of U.S. Corporate Taxation on the U.S. Economy.

Determining the effect corporate taxation might have on the economy is a complex and challenging undertaking. What Martin's (1991) and King's (1993) analyses make clear is that policymakers have *intended* that corporate tax policy stimulate investment and growth. Sørensen (1995, 280) shows big

[4] See the Committee on Ways and Means 1993, 78-81 on the "two-tier" taxation of corporate income.

differences among various taxation theories regarding their predictions about the effects of the "classical" (U.S.) corporate tax systems on economic activity. (See also Engen and Skinner 1996.) In light of these review, we expect that the structure of the relationship between changes in corporate taxation and changes in economic variables depends in part on the initial levels of taxation. That is, changes in higher levels of taxation will differ in effects from changes in lower levels of taxation. In the 1981-98 period studied here, corporate tax revenues as a percentage of GDP ranged from 1.8 to 3.2%, levels far lower than the 5 to 6% range common in the 1950s. We expect that these comparatively low levels are less deleterious to investment and growth than higher levels might be, and that modest changes in tax revenue or rates have little short-term economic impact.

Corporate Taxation and Partisan Politics.

The conventional view of the Democratic and Republican parties regarding corporate taxation is that the parties differ on their policy preference (Edsall 1988). Democrats are highly likely to pursue economic growth strategies consistent with increasing corporate taxes relative to other sources of revenue, whereas Republicans are highly likely to pursue economic growth strategies consistent with decreasing corporate taxation (Hibbs 1987, Quinn and Shapiro 1991b). These partisan effects are short-run as Administrations will not continually increase (or decrease) corporate tax burdens, but are instead likely to shift corporate tax collections to a preferred level, and maintain that level.

In contrast, Williams and Collins (1997) reject the partisanship thesis on theoretical grounds because partisan differences arising from change in the composition of the government is another source of time-inconsistency. The implication is that partisan differences should not matter much. Williams and Collins found that various partisanship measures were not a statistically significant force in their examination of tax policy (1997, 230, fn. 25).

Dryzek (1996) develops what he suggests are the political implications of open capital markets. He claims that democratic choices and institutions are being eroded by the global capitalist economy, and that the "operative mechanisms" that constrain democratic choice are "purely automatic" (1996, 17-34). Policies adverse to business produce capital flight and capital strikes, and "astute government officials will therefore anticipate and attend to the reactions of markets" (1996, 26). The implication of this version of the "structural dependence of the state on capital" theory is that partisan politics and electoral choices should matter little regarding taxation of business firms. (See also Lindblom 1977.)

The Effects of Business PAC Activities.

A core finding of previous work on corporate taxation is that business political activity is an important influence. (See Bowles, Gordon, and Weisskopf 1989, Jacobs 1988, Quinn and Shapiro 1991a.) Williams and Collins (1997) hypothesize that business political activities joined the list of variables that are not an influence on corporate tax policy. They find (1997, 230-32) that an increase in business political activity leads to *increased* corporate taxation. Increases in the number of corporate political action committees also leads to *increasing* GDP, as suggested by the results of their VAR analysis (Williams and Collins 1997, 232). Political activity by businesses, then, possibly contributes to economic growth. This finding is surprising to say the least. Our expectation is that once cointegrated relationships and omitted variables are accounted for, the economy-wide result of business political activity is unlikely to be positive.

The Origins of Business PAC Activities.

Why do firms mobilize politically? Previously, scholars have suggested that business mobilization occurs in response to poor economic conditions, among other things. The election of a Democratic president has been found to precipitate increased business political activity; business leaders, especially of small to medium size businesses, have traditionally seen Democratic administrations as less favorable to their interests than Republican ones. (See Eismer and Pollack 1988.)

A failing of many previous studies of corporate taxation is a lack of attention to the dynamics of the relationship between taxation and business PAC activity, an omission explicitly addressed by Williams and

Collins 1997. They find empirical support for the hypothesis that increases in corporate taxation leads to increases in business political activity.

Hypotheses.

We propose the following hypotheses:

1) Corporate taxation is influenced by economic events. An increase in GDP leads to a short-term increase in corporate taxation revenue. International economic variables also influence corporate taxation revenue: increased goods and service exports and capital imports raise short-term tax receipts.
2) Democratic presidents differ from Republican presidents in their corporate taxation policies, with Democratic administrations being associated with higher short-term tax revenue.
3) An increase in the numbers of business political action committees (compared to other actors) leads to lower short-term corporate tax revenue.
4) Business interests will mobilize by increasing the numbers of business PACs in response to Democratic Administrations and increasing corporate taxation.

MODELS and METHODS

Vector Autoregressive Models.

We model the relationships among taxation and the other economic and political variables using a vector autoregressive (VAR) model of order p:

$$Y_t = A_1 Y_{t-1} + A_2 Y_{t-2} + \ldots + A_p Y_{t-p} + \Phi D_t + \varepsilon_t \qquad t = 1,2,\ldots,T$$

where Y_t represents the *k*x*1* vector of interest, $Y_t' = (Y_{1t}, Y_{2t}, \ldots, Y_{kt})$, the deterministic components are represented by D_t (intercept and possible time trend or seasonal dummy variables), and ε_t is the error term assumed to be independent and identically distributed $N(0, \Omega)$. VAR models capture the dynamic relationships among all the variables under study without having to specify which among them are exogenous and endogenous. VARs usefully allow for various lags on those effects and are especially useful for examining possible feedback effects.

Our base model of corporate tax revenue includes a measure of corporate tax liability as a percentage of national income (Tax/GDP) or assets (Tax/Assets). As variables in the analyses, we also use gross domestic product, investment, foreign direct investment inflows and outflows, international trade flows, a measure of corporate political activity, and various measures of partisanship.

In preliminary graphical exploration of the data, we saw indications that some variables exhibited nonstationary behavior. The model might, therefore, have some unit roots and cointegrating relationships among the variables. To confirm this first impression, we tested for unit roots in each of the series using Augmented Dickey-Fuller tests (see Appendix and Table A-2) and tested for cointegration using the Johansen (1995) likelihood-based methodology. We then estimated the model in vector error correction (VEC) form, and we summarized the impact of each variable on the others using impulse-response functions.

Vector Error Correction Models.

We did not estimate VAR models for levels of the variables because our models include some nonstationary series. When impulse-response functions are estimated from an unrestricted (level) VAR model, the existence of unit roots makes the estimates of the impulse response functions inconsistent at long horizons, and their small sample distributions show substantial variability in Monte Carlo experiments.[5] A solution is to estimate, instead, the model in error correction form (Phillips 1998, 30).

[5] Further, the asymptotic distribution of the estimated coefficients of a VAR model that has unit roots is different from the asymptotic Normal distribution that characterizes the estimated coefficients (and impulse response functions) of a stationary VAR model. The estimators of some of the parameters converge at a faster rate (are hyperconsistent), and the asymptotic Normal distribution has a singular variance-covariance matrix. The impulse response functions from VAR with unit roots do not go to zero as the lead-time increases, but carry the effect of the unit roots indefinitely.

In very simple terms, a VEC model is a VAR model in differences with some extra variables, the error correction terms, which are stationary linear combinations of all the level variables in the model. To be precise, let ΔY_t be the first difference of Y_t, given by $\Delta Y_t = Y_t - Y_{t-1}$. In the Johansen (1995) approach, the VEC model can be written as

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \ldots + \Gamma_{p-1} \Delta Y_{t-p+1} + \Phi D_t + \varepsilon_t$$

where $\beta' Y_{t-1}$ represents the cointegration relationships, which are stationary. Here, $\beta' Y_{t-1} = 0$ are interpreted as long-term relationships. If these relationships do not hold exactly at some point in time, there is an equilibrium error $\beta' Y_{t-1}$, and α measures the speed of adjustment. This model is also known as Reduced Rank Regression (RRR) because the matrix of coefficients of the lagged levels variable (Y_{t-1}) takes the form $\alpha\beta'$, where α and β are kxr matrices of rank $r<k$. There are r cointegration relationships and k-r unit roots in the model. Estimating the reduced rank regression means that we do not have to estimate unit roots. It is very important that both the number of lags in the model and the cointegration rank are correctly specified, so we need to know how many unit roots and how many cointegrated relationships there are in the model.[6]

Considering all these issues we took the following approach. First, we conducted unit-root tests (described in the appendix) and then used Johansen's tests to determine if there are cointegration relationships. GDP and Investment are I(1) variables and are cointegrated for the 1981-1998 period.

Next, we estimated the VEC model. Note that when there are some I(0) and some I(1) variables in the model, which is true in this case, the matrix β must include what Lutkepohl (1999) terms a "trivial" cointegration equation for each I(0) variable in the model. Hence, the error correction term reflects not only the cointegration between GDP and investment, but also the trivial equations.

The model was estimated using Johansen's maximum likelihood approach using Hansen and Juselius's (1995) CATS software. From the resulting VEC model, we obtained impulse-response functions.

Impulse-Response Functions From a Vector Error Correction Model.

Impulse-response functions show how a shock to one variable is followed by changes in the other variables over several periods of time, once the effects of all variables in the model are accounted for. In the case of a stationary VAR, the model can be written in vector MA(∞) form

$$Y_t - \mu = \varepsilon_t + \Psi_1 \varepsilon_{t-1} + \Psi_2 \varepsilon_{t-2} + \ldots$$

and each matrix Ψ_s is

$$\frac{\partial Y_{t+s}}{\partial \varepsilon'_t} = \Psi_s$$

so the (i,j) element of Ψ_s measures the consequence of a one-unit change in the innovation of the j-th variable at time t on the i-th variable at time $t+s$, holding all the other innovations at all dates constant.

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[6] Another possibility would have been to use fully modified vector autoregression (FM-VAR) methods, which have some highly desirable properties (Phillips 1995). See Freeman et al. 1998 for a discussion.

A consideration in choosing between VEC and FM-VAR is that none of the available programs for either procedure contains a method of generating reliable boundaries for impulse-response functions, which are the main tool for estimating influences. In this investigation, the development of a method for generating confidence intervals is one of the contributions of the article.

We chose to develop a procedure for establishing confidence intervals for the impulse-response functions of VEC models because, as reported in Freeman et al. (1998, 1312), with sample size of 50, the tests based on FM-VAR reject a false null hypothesis only between 7% and 23% of the time. With a sample size of 100, tests based on FM-VAR reject a false null hypothesis 9% to 49% of the time. To our knowledge, no similar study has been carried out comparing FM-VAR to the Johansen approach. Our sample has 71 time periods, so what we know is that the FM-VAR methods are much more likely than not to produce misleading results.

A VAR process with unit roots does not have an MA representation. Even so, the Ψ_s matrices can be obtained using the same recursive relationships as in the stationary case:

$\Psi_o = I$, and $\Psi_i = \Sigma^i_{j=1} \Psi_{i-j} A_j \quad i = 1,2,\ldots$

As in the case of stationary VAR, the investigator must obtain the impulse-response to orthogonal shocks. That means we need to obtain the Choleski decomposition of the error variance-covariance matrix Ω. We use the lower triangular matrix P, such that $PP' = \Omega$. Thus, the impulse-response to orthogonal shocks is given by $\Theta_i = \Psi_i P$, i = 1,2,… The main difference from the impulse-responses for stationary VAR models is that in the case of systems with unit roots, the effect of a one-time shock may not die out asymptotically.

The VEC model does not produce estimates of the A_j matrices directly. It is necessary instead to obtain them from the short-term matrices (Γ_j, j = 1,…,p-1) and the error correction coefficients (α and β). The resulting Ψ_s and Θ_i matrices are the same as those obtained by estimating a VAR model in differences with the extra term $\alpha\beta' Y_{t-1}$, and then obtaining the impulse-response function for the original variables in levels.

Typically, impulse-response functions are presented in graphical form, with one graph for each (i,j) element of Ψ_s, s = 0,1,… Those graphs present estimates of the impulse-response function, but it is also necessary to provide some information about the uncertainty of those estimates. The assessment of the variability of the estimated impulse-responses is presented in the form of confidence intervals or bounds. When the confidence intervals surrounding the estimated impulse-response function do not include zero, the estimated response is said to be statistically significant.

We obtained the boundaries presented in the graphs using the empirical 5^{th} and 95^{th} percentiles from a Monte Carlo simulation of the model with the estimated parameters. It is important to use percentiles when obtaining these confidence intervals because the sampling distribution of the estimated impulse-response functions is skewed, particularly in models with unit roots. The calculations required to obtain the boundaries for the impulse-response function in a VEC model are not available in commercial econometric software, but are available upon request. (See the Appendix for a description of the procedures we followed.)

When interpreting the impulse-response results, we must keep in mind that the order of the variables in the model matters in the following way: the contemporaneous correlation is assigned as a response to the variable that appears first in the model. The first variable is the only one with a potential immediate impact on all other variables. The second variable may have an immediate impact on the variables that follow, but not on the first one, and so on. This feature is a result of the choice of orthogonalization method, and the order of the variables is a restriction imposed on the model. In this analysis, we report the results where GDP is the first variable and corporate taxation was treated as a contemporary response to the economic variables.

We do not report the VEC coefficients because they cannot be interpreted as regression coefficients. A one-time-unit innovation in one variable may affect several other variables, and their cumulative impacts have to be worked out through the system. This is why the impulse response-functions give a better picture of the relationships among the variables. Neither do we report the cointegration relation coefficients per se because, according to Juselius, "… cointegration between variables is a statistical property of the data that only exceptionally can be given a direct interpretation as an economic steady-state relation" (1999, 25).

DATA AND MEASURES

The complete data were available on a regular basis only from the first quarter of 1981 to the third quarter of 1998, yielding seventy-one quarterly observations. These quarters included the first and second Reagan Administrations, the Bush Presidency, the first Clinton Administration, plus seven quarters of the second Clinton Administration. We used data from the last three quarters of the Carter Administration as information about initial conditions: i.e., 1980:4 is t=0; 1980:3 is t=-1, and 1980:2 is t=-2. Events of the Carter Administration were otherwise unrepresented in this analysis.

Corporate Taxation Variables.

By themselves, statutory corporate tax rates are not meaningful in assessing the taxes on businesses because the tax code contains almost innumerable profit income exclusions. For example, the statutory tax

rate for corporate income over $100,000 was 46% under the 1982 Tax Equity and Fiscal Responsibility Act (TEFRA), the first "Reagan" tax act. This compares to the current tax rate of 35%-39% for corporate income over $100,000. The higher TEFRA rates notwithstanding, corporate tax payments as a percentage of GDP were 1.8-2.0% under TEFRA, compared to 2.9-3.2% of GDP under the current law. The lower tax revenue to the Treasury from TEFRA came from TEFRA's exceptionally generous income exclusions.

What we can observe directly are the revenues collected from corporate taxpayers, and we use these observed tax receipts in this analysis. To account for the origins and consequences of corporate taxation, we used data for all corporate businesses, financial and nonfinancial. The National Income and Product Accounts (NIPA) of the U.S. Department of Commerce provided quarterly data on corporate tax payments. Specifically, we used NIPA profit tax liability data as the numerator for corporate taxation.

But, what is the appropriate benchmark for assessing tax receipts over time? Time-series data on tax liabilities as such are not very meaningful without controlling for either inflation or increases in economic activity. We considered five denominators to account for inflation and growth: gross domestic product, the capital stock of corporate businesses, corporate earnings before interest and taxes (EBIT), corporate earnings before interest, taxes, depreciation, and amortization (EBITDA), and taxable corporate income with inventory valuation adjustment (IVA) and capital consumption adjustment (CCA).

Each denominator has advantages in attempting to assess our hypotheses. In using gross domestic product, we were able to compare tax revenue against the level of overall economic activity, while controlling for the effects of inflation. These data are publicly available and are used frequently as a common benchmark in political economic analysis. Corporate tax revenue as a percentage of GDP is one of two comparative indicators used by the Committee on Ways and Means in their statistical descriptions of U.S. corporate tax policy (see, e.g., 1993, 85-86 and 254-6). We see no disadvantages in its use.

Corporate capital stock is a measure of the accumulated value of corporate productive capacity. Many economists believe that tax over capital stock or some other measure of corporate assets is perhaps *the* appropriate measure for evaluating the tax burden of corporate businesses over time. The disadvantage in this analysis of quarterly data is that estimates of U.S. corporate capital stock are available annually, so some interpolation of annual data was necessary to establish a quarterly series.

Another approach is to use an accounting measure of corporate profits (i.e., EBIT or EBITDA).[7] EBIT and EBITDA reflect what income from continuing operations is available to the firm, prior to any of the tax and interest considerations. EBITDA has the further advantage of excluding from consideration the various schemes for calculating depreciation. Because both EBIT and EBITDA are indicators of the cash flow available to the firm, both indicators are widely used by stock analysts and shareholders in evaluating the performance of managers of large U.S. corporations. Comparing annualized values of Tax/EBIT and Tax/EBITDA from the National Income and Product Accounts to the Mendoza, Razin, and Tesar (1994, 309) estimates of effective U.S. corporate tax rates yields correlations of .95 and .96, respectively.

A disadvantage of using these measures is that managers have incentives to under-report earnings in some quarters, and over-report earnings in other quarters. Indeed, the Securities and Exchange Commission has suggested that the effects of "managed earnings" on corporate balance sheets have grown large, and are potentially very distorting. (See, e.g., Levitt 1998, Walker 1999.) In this investigation, the resulting "smoothing" and "puffing" of corporate profits produces either decreased variance in the dependent variable or a trend, respectively. Decreased variance induces downward bias in the estimates of the influences of other variables, which potentially undermines the validity of the investigation.

A third approach, followed by Williams and Collins 1997, is to use corporate taxable income as a denominator. The advantage is that the IRS uses profits to determine taxation, and this measure can therefore be considered an indicator of corporate tax rates. The severe disadvantage of using a measure of tax receipts over taxable income is that *cuts* in corporate taxable income (the denominator) will sometimes

[7] EBIT is line 4 minus line 5 in Table 1.16. EBITDA is line 4 minus line 5 plus line 2 (depreciation) in Table 1.16.

appear to be *increases in these tax rate measures.* This occurs because tax cuts are sometimes brought about, not by changing the rate of taxation, but by changing what counts as taxable income.

Accelerated depreciation provisions, for example, reduce the taxes firms pay. The Accelerated Cost Recovery System (ACRS) from the 1981 Tax Act (section 168 of the code) was advantageous to many firms. ACRS cut the income that counted as taxable income; the denominator (taxable income) used to establish the effective corporate tax rate. Thus, income exclusions, while actually reducing taxes, will wrongly appear to be tax *increases* because the denominator (profits) decreased proportionally more than the numerator (tax liability), particularly in the presence of Alternative Minimum Tax provisions. (See Lyon 1997.) The reverse is also sometimes true: the 1986 Tax Act modified section 168 and increased the taxes firms paid when ACRS was replaced with MACRS (the Modified Accelerated Cost Recovery System). But, because income is added to the denominator, the actual magnitude of the tax increase is obscured with ECTR.[8]

The denominator used to provide a benchmark for corporate profit taxation should not distort the results of empirical analyses. We therefore use economy-wide measures for the denominator. To increase confidence in the robustness of our results, we used both Tax/GDP and Tax/Assets in our analyses. Our view is that, in a multivariate dynamic setting, changes in corporate tax receipt data over GDP or capital stock offer reliable indicators of U.S. tax policy, as economic forces beyond the control of the government can be accounted for. For comparison purposes, we also report results using the effective tax-rate indicators, though we regard them with more caution because of the earnings management problem. We do not use Tax/ECTR because that analysis will produce misleading results.

Figure 1 shows the quarterly tax liabilities of U.S. corporate businesses, 1953 to 1998, as a percentage of either U.S. gross domestic product (GDP) or its corporate capital stock. Figure 2 shows the same numerator, quarterly corporate tax liabilities, as a percentage of two indicators of either EBIT or EBITDA. From Figures 1 and 2, we see a gradual downward trend from the 1950s through the late 1970s in U.S. corporate taxation. The graphs of the data in Figure 1 show an upward trend from 1981 on. The Tax/EBIT series, in contrast to the others, moves downward in the last few years of the 1990s, despite the absence of changes in tax law. The disparity between Tax/GDP and Tax/Assets versus Tax/EBIT in the 1990s might arise from the earnings "puffery" described in Levitt (1998) and Walker (1999).

[Figures 1 and 2 about here]

Business PAC Activity.

Business political activity is difficult to measure and isolate in the aggregate. As Schuler (1999) shows, whether or not a firm employs a PAC is highly correlated with other types of corporate political activity (e.g., whether the firm has a Washington office, undertakes lobbying activity, has a membership in a trade association). We use the number of business political action committees relative to other political action committees as a proxy for the relative political activity of businesses, as in Quinn and Shapiro 1991a and Williams and Collins 1997.

The Federal Election Commission (FEC) has released PAC data on a semi-annual basis since 1979, and it has internal data available on a quarterly basis from the third quarter of 1981 on. Joining these data series produces reliable data from the first quarter of 1981 on.[9] Prior to 1979, the data were released annually only. One additional consideration is that corporate PACs were not legal before 1971 (Epstein 1969). Prior to November 24, 1975, which was the date of the FEC Advisory Opinion "SUNPAC," their legality was

[8] Another consideration is that corporate profits with IVA and CCA contain information about substantial profits from foreign direct investment. (See Table 6.16 of the NIPA.) It is important, therefore, to account for capital flows in the analysis when using ECTR.
Another concern in using Tax/ECTR is that the most recent observations of the denominator are provisional estimates of taxable income taken from the financial reports, not tax returns, of U.S. corporations. (Phone interview, Dennis Quinn and Ken Petrick, Department of Commerce, 31 March 1999.) The ECTR data were fully reliable only prior to 1996.

[9] The semi-annual data are available electronically at www. Fec.gov/press/paccount.htm. We used the data for three prior quarters to 1981(1) for use in the initial conditions. We interpolate linearly the PAC data as the arithmetic mean of the data on either side of the missing observation.

highly questionable, whereas other PACs, such as trade union PACs, have existed since the 1940s. Consequently, the corporate PAC data as a percentage of all PACs have a "spike" from 1976 through 1979 as firms began to organize PACs at a rate faster than that for other organizations, with the high point of the "spike" occurring during the midyears of the Carter Presidency. While it is possible that the mobilization of corporate PACs was a response to the policies of President Carter, it is also possible that business PACs organized quickly during this period because they were legally permitted to do so for the first time. In estimating the models from 1981(1) on, we can see whether the divergent estimates of business PAC effects in Quinn and Shapiro 1991a and Williams and Collins 1997 are present if the business PAC "spike" is excluded from analysis. Figure 3 graphs available Federal Election Commission observations for PAC data.

[Figure 3 about here]

An important empirical issue is the choice about which period to attribute the PAC data. The FEC data are "point-in-time" data.[10] We treat the PAC data as being "effective" from the release date. For example, the FEC release of December 31, 1991 would be attributed in our analysis to the first quarter of 1992. We do so because point-in-time data are precisely that: the state of the world as announced then and as it prevails for a while. Williams and Collins (1997) make a very different choice, though note that quarterly data were unavailable to them. They attributed the data released by the FEC on 31 December 1991, for example, to the period commencing on 1 July 1991 and ending with 31 December 1991.[11] The FEC data released on 1 July 1991, in turn, was attributed to the period 1 January 1991 to 30 June 1991, and so on.[12] A six-month difference in treatment can matter in a quarterly based time-series investigation employing VEC methods.

Partisanship.

Partisanship was operationalized by using a dummy variable for the party of the President: 1 for Democrat, 0 for Republican. We used the party of the President to represent partisanship because of the many studies showing the dominance of the President over Congress on tax legislation. We also estimated a separate model with the Presidential partisanship variable and a separate variable representing partisan Congressional influence, which was measured by the proportion of the seats controlled by the dominant party in each branch.[13] A consideration is that any partisanship indicator is binary (as in the case of "presidency"), which violates the assumption of normality. The U.S. has only two major political parties, however, and omitting estimates of partisanship on taxation is a worse problem.

To ensure the robustness of the results, we also estimated models with a partisanship indicator that included information about joint partisan control of the Presidency and Congress. For this indicator (President&Congress%), we use percentage of seats as the indicator of partisan control of Congress, with a positive sign if Democrats are majority and a negative sign if Republicans are majority, and added it to an indicator for partisan control of the Presidency (1=Democrat, -1=Republican).[14] The results on Presidential partisanship were highly robust.

Economic Variables.

[10] Phone interview, Dennis Quinn with Dr. Robert Biersack of the Federal Election Commission, September 4, 1997.

[11] Email correspondence, Brian Collins to Dennis Quinn, September 26, 1997.

[12] Williams and Collins (1997) employed an interpolation method as actual quarterly data was unavailable at the time of their study. The interpolation routine is available from Professor Williams at jotwilli@indiana.edu.

[13] For example, the Senate Republicans had 55% of the Senate seats in the 105th Congress, 1996-1998, and this is scored as -.55. House Republicans held 52.3% of the seats in the 105th Congress, which is scored as -.523. Averaging these numbers together gives us a partisanship score of -.536 for the 105th Congress.

[14] For example, the partisanship score of the 105th Congress is -.536. This indicator was added to the indicator for Democratic control of the Presidency (i.e., 1), for a combined score of .463. This indicator gives higher weight to partisan control of the Presidency compared to partisan control of Congress, which is appropriate given the literature on Presidential dominance on corporate tax policy.

The domestic and international economic variables were taken from the Department of Commerce's National Income and Product Account (NIPA) Tables, from the Federal Reserve's *Balance Sheets, Historical Data* (Z.1), and from the International Monetary Fund's *International Financial Statistics*. Corporate tax liabilities (Table 1.14 or Table 1.16), investment (Table 5.2), gross domestic product (Table 1.1), and trade (Table 4.2) were obtained from the first source. Foreign Direct Investment (FDI) came from the IMF's IFS, and corporate capital stock (as tangible assets)[15] was taken from the Federal Reserve. These data were also available through the third quarter of 1998. The data for corporate capital stock are available on an annual basis only, and a linear interpolation is used to establish a quarterly series.

The taxation, trade, foreign direct investment, and investment variables used are ratios, which put the variables on comparable terms in the same time period, and controlled for the effects of inflation. The investment variable consisted of each quarter's plant and durable equipment investment as a percentage of last quarter's tangible assets. FDI and trade variables were treated as percentages of gross domestic product. Gross domestic product, when used as a separate variable, was specified in real terms (1992 chained dollars) and logged for standard econometric reasons (e.g., better-behaved residuals).

For international economic effects in our model, we used trade flows as proportions of GDP, which has been a standard measure used in estimating the effects of international trade openness on corporate taxation, as in Swank 1998. We report models with FDI operationalized with separate FDI inflows and FDI outflows indicators because the variables have a correlation of .16 in our sample, and exhibit differing dynamics. We also report some results from models where capital flows are operationalized as the net flows of FDI. These results were similar to those of models where inflows and outflows are used.

One empirical consideration about estimating the relationship of international economic forces and corporate taxation is that capital and trade flows are closely connected. Tang (1997) reports estimates from the U.S. Department of Commerce that 40% of world trade and 38% of U.S. trade is intrafirm trade, so that FDI flows are likely to lead to greater trade flows. Moreover, in the system of national accounts used to record exchange among nations, trade and current account surpluses (deficits) are necessarily offset by more or less equivalent capital account deficits (surpluses), plus or minus errors and omissions. The effects of international economics variables are likely to be understated when components of the capital account and the current account are both represented in the model, as is true here. This bias is especially pronounced as we have comparatively few degrees of freedom, and any additional variables are an added burden to the efficiency of the estimates. To omit either trade or capital indicators, however, is to risk model misspecification, which is a worse problem, so we include them.

RESULTS

Unit Roots and Cointegrated Relationships.

When building a VAR model, investigators need to understand the univariate properties of the time series involved and whether or not the series have unit roots. Appendix Tables A-1 and A-2 give a summary of our findings. The conclusion from the univariate analysis is that only some series have unit roots.

Do the tax liabilities data from 1954 to 1980 have similar statistical properties to the data from 1981 to 1998? No. From 1954 to 1980, we found that three of the four tax series had unit roots: that is, they are I(1). The fourth series, Tax/EBITDA, had a downward negative trend.[16] Williams and Collins (1997, 226) note that ECTR also has a unit root and a negative deterministic trend from 1953-94. In contrast, in the 1981-98 period, we found that measures of both corporate tax revenue and effective corporate tax rates were without unit roots (that is, are I(0)). Further, the unit root tests also revealed that, in the latter period, Tax/GDP, Tax/Assets, and Tax/EBITDA had *upward*, not downward, deterministic trends (see Table A-2).

Based on this evidence, U.S. corporate tax receipts 1954-80 had a long-run relationship with GDP and investment. These results are fully expected, conforming to the standard view of the U.S. corporate tax

[15] The tangible assets data come from "btabx.zip" file from http://www.federalreserve.gov/releases/Z1/data.htm.

[16] Furthermore, the two corporate tax revenue measures from Figure 1 are cointegrated with two other I(1) series, gross domestic product and investment.

code from Presidents Eisenhower through Carter, which is that the code enabled firms to reduce taxes via higher levels of business investment, which then produced higher rates of growth.

Gross Domestic Product (i.e., Logarithm of real GDP or LRGDP), Investment (i.e., Producer's Durable Equipment as a percentage of lagged Tangible Assets or PDE/Assets), corporate political action committees as a percentage of all PACs (CRATIO), and our indicator of partisan control of Congress are all I(1). Foreign direct investment outflows (FDIout) join the tax variables as I(0).

In the VEC models we estimated, GDP and Investment are cointegrated and are part of a long-run structural relationship. The other variables in the analysis do not exhibit long-run relationships, but are possibly endogenous in the analysis to other variables. These short-run relationships, if present, were captured in the impulse response function.

Impulse-Response Functions.

Figure 4 presents a partial set of impulse-response functions, restricting our attention to those graphs where a tax variable is represented. In Figure 4, the VEC model used contains the following eight variables: TAX/Assets (Figure 4a) or TAX/GDP (Figure 4b), LRGDP, Invest=PDE/Assets, FDIin/GDP, FDIout/GDP, TRADEtotal/GDP, CRATIO=Corporate PACs/All PACs and Pres=Presidency. We report the impulse-response results from the contemporaneous periods and the subsequent eight quarters. The boundaries use the 5^{th} and 95^{th} percentiles of the impulse-response function at each lead time, generated through Monte Carlo methods. Our attention was focused on results consistent across models.

[Figure 4 here]

International Economic Variables. To which of these variables, then, did corporate tax revenue respond? In Figure 4, both indicators of corporate tax revenue are statistically significant and positive responses to inward Foreign Direct Investment, though the duration of the estimated effects is greater in the model with TAX/GDP. Both indicators responded positively and statistically significantly at lags 0 and 1 to TRADE.

From Figure 4, we see that trade was a consistent response to changes in U.S. corporate tax revenue (positive at lag 1). The statistically significant responses of FDIin (positive) and FDIout (negative) that we see in the models with TAX/GDP are not found in models with TAX/Assets.

The positive response of both corporate tax revenue indicators to FDI inflows is of special interest, in light of the debate about whether or not affiliates of foreign-owned enterprises paid appropriate levels of U.S. corporate taxation. The consistency of the positive response to TRADE is also of note, given prior findings.

When the influence of FDI was operationalized as net FDI (FDIin-FDIout), both tax indicators responded positively to FDInet. FDInet, in turn, was always a *positive* and strong response to increased taxation in all models we estimate. From the perspective of structural dependence theory, this "wrong" sign is a puzzle as it implies that increased taxation led to increased net FDI inflows.

What we can safely say is that internationalization seems to have led to increasing tax revenue to the U.S. government. Increasing trade and increasing FDI inflows led to higher corporate tax revenue. In turn, we saw no systematic evidence that corporate taxation diminishes trade or FDI inflows in the U.S. Maintaining high revenue yields from corporate taxation while maintaining international economic openness is a feat mastered by the U.S.

The U.S. Domestic Economy. In Figure 4, we see that corporate tax revenue always responded to change in GDP in a positive and statistically significant fashion. Corporate taxation was a response to domestic as well as international economic activity. Neither revenue measure responded to investment.

The U.S. gross domestic product was a negative response to increasing tax in Figure 4, although this effect depended on GDP being entered as the initial variable in the VEC. The effects in Figure 4 occurred with a one-quarter lag and were very small. Investment was not a consistent response to tax revenues.

Partisanship. Partisanship matters: Figure 4 shows that corporate tax revenue responded to partisanship, with Democratic presidents leading to higher corporate taxes, and the election of Republican presidents leading to

lower corporate taxes. This finding was extremely robust in all models estimated with a partisanship dummy and is consistent with and Quinn and Shapiro 1991a, but not Williams and Collins 1997.

When we added an indicator of partisan control of Congress to the models in Figures 4 and 5, the indicator for partisan control of Congress was not statistically significant. The estimated influence of the measure of joint partisan control of the Executive and Legislative Branches was very similar to that for the Presidential dummy variable alone.

Corporate taxation is not the only response to partisan change. Corporate PACs as a percentage of all PACs was a strong and statistically significant response in all models to partisan change in the executive and legislative branches, with Democratic control being associated with marked increases in business PAC activity (see Table 1). This finding is consistent with Eismer and Pollack's (1988) study. Note that these results do not include information about the mobilization of business PACs in the four years following the SUN-PAC decision in November 1975, which were also the main years of the Carter Presidency. Hence, the mobilization of business interests, 1976-79, did not influence our results.

Another consistent result is that FDI outflows were a small, but positive and statistically significant response to presidential partisanship. The election of a Democratic (Republican) president was associated with a subsequent increase (decrease) in FDI outflows.

Business Political Action Committees. How did corporate tax revenue respond to business political activity? Table 1 summarizes the results of the impulse response functions for Business PACs. We found that both TAX/GDP and TAX/Assets were negative responses to increasing corporate PACs. A relative increase in corporate activity (PACs) resulted in a decrease in government tax revenue, which supports the proposition that business political activity is a determinant of corporate tax revenue. The results are not influenced by switching the order of these two variables and are robust to shifts in the timing of the PAC variable.

[Table 1 about here]

How do the results from the effective tax rate models compare? Figure 5 shows the VEC impulse response functions for the same base model as in Figure 4, except the tax variables are tax rates: TAX/EBIT (Figure 5a) or TAX/EBITDA (Figure 5b).

We found that both TAX/EBIT and TAX/EBITDA show positive, statistically significant contemporaneous responses to both TRADE and FDIin. Neither tax rate indicator was a response to FDIout. As in the models with the revenue indicators, TRADE was a positive response to TAX/EBIT and TAX/EBITDA. FDIin (positively) and FDIout (negatively) were responses to increases in corporate tax rates. Note that the estimated negative effects of tax rates on FDI outflows are half the size of the estimated positive effects from FDI inflows. As with Figure 4, we see that corporate tax rates always respond to change in GDP in a positive and statistically significant fashion. Neither tax rate indicators are consistent responses to investment.

The U.S. gross domestic product is a negative response to increasing tax rates in Figure 5, though this effect depended on GDP being entered as the initial variable in the VEC. Investment is a negative response to both corporate tax rate indicators.

Figure 5 shows that corporate tax revenue and corporate tax rates respond to partisanship, with Democratic presidents leading to higher corporate taxes, and the election of Republican presidents leading to lower corporate taxes. This finding was extremely robust in all models estimated.

CONCLUSION

Our results can be summarized succinctly. As shown in this study of 1981-98, international economic openness in capital and trade leads to increasing corporate tax revenue collections. Economic growth also leads to increased corporate taxation. Corporate taxation, both revenue and rates, is a response to "normal" politics: Democratic Administrations increase corporate taxation, and Republican Administrations decrease it. Corporate tax revenue is a negative response to the numbers of corporate PACs relative to other PACs. The percentage of corporate PACs is a positive response to the Democratic Presidential Administrations and to increases in the effective corporate tax rate. Neither increased tax

revenue nor rates are a detriment to increasing trade and foreign direct investment inflows, and are at most only a modest stimulant to increased FDI outflows.

A question that arises is how the VEC models used here might help us understand longer-run relationships. We focused on the short-term relationships, 1981-98, because the tax variables in our analysis are characterized by endogeniety and stability and by the absence of long-run relations. In the 1954-80 period, however, corporate tax receipts were I(1) and cointegrated with GDP and investment. Further analyses of the cointegration relationships from the prior period could allow for the discovery of the sign of those relationships and, potentially, the determination of the approximate dates of the structural break in the cointegrated relationships. This approach could provide greater insight into the welfare consequences of these prior policies.

To conclude, we take special note of the incongruence between these results and what might have been expected from structural-dependence theory. An expectation from structural-dependence theory is that international capital and trade openness leads to diminished state capacity to tax corporate income (Dryzek 1996). Another expectation is the absence of the effects of normal politics on corporate taxation (Williams and Collins 1997). None of these expectations are consistent with what we find in our analyses of U.S. corporate taxation in the 1981-1998 period. The evidence from this study of one aspect of government-business relations in the U.S. is instead consistent with a "normal" politics view of political economics. Who gets what when depends on (a) who wins what election, (b) who organizes, and (c) how economies here and abroad are doing.

APPENDIX

Unit Roots. The challenges of determining whether a series has a unit root are as follows. First, the null hypothesis in the test is that the series has a unit root. That hypothesis is "maintained" until the evidence in the data shows it should be rejected, but that does not mean we have evidence to support the null hypothesis that there is a unit root. Second, these tests are sensitive to the inclusion or exclusion of the deterministic components of the series, which are the intercept and a time trend. Assessing the series by itself, the series with the time trend, the series with the intercept, and the series with both intercept and time trend require different tests for each analysis, and each test has its own distribution, and thus its own special critical values. Third, if the series presents a structural change because of a level shift, the test may wrongly indicate there is a unit root. A level shift in a series has different implications from the presence of a unit root, so further testing is needed (Campos, Ericsson, and Hendry 1996). Fourth, the unit root tests have low power (Maddala and Kim 1998, chapter 4). Finally, after undertaking this long sequence of tests, each of which is conditional on the results of the previous tests, it is impossible to determine the significance level of the whole process. Still, it helps the analysis to make an assessment of the univariate properties of the series, so we obtained autocorrelation functions, tests for unit roots and autoregressive models for each of the series under study.

Autocorrelation Functions. We considered first the pattern in the autocorrelation function (ACF). A slow decay in the ACF together with one large spike at lag 1 in the partial autocorrelation function usually indicates that the series is nonstationary. (This pattern also appears when there is a level shift in the series.) On the other hand, a fast decay in the ACF indicates the series is stationary.

Augmented Dickey-Fuller (ADF) tests for unit roots. These tests require estimating a regression where the dependent variable is the difference of the variable of interest (for example, ΔY_{1t}), and the explanatory variables are the level of the variable of interest (Y_{1t}) and several lagged values of the difference ($\Delta Y_{1t-j,}$ j=1,2,…,p), as in the following equation:

$$\Delta Y_{1t} = a_0 + \gamma Y_{1t-1} + a_2 t + \Sigma^{p}_{j=1} \beta_j \Delta Y_{1t-j}$$

In order to improve the power of the tests, we use the procedure developed in Enders 1995. We first obtained the ADF test including trend and intercept, and as many lags of ΔY_{1t} as needed, starting with a large number of lags (eight in our case) and eliminating those that are not significant. We indicated the number of lags used in each test in parenthesis next to the value of the ADF test reported in Tables A-1 and A-2. If the null hypothesis of unit root was not rejected, we tested whether a trend term is needed (test $a_2=0$ given $\gamma=0$). If the trend is not significant ($a_2=0$), we repeated the ADF with the intercept only, and as many lags of ΔY_{1t} as needed. When the trend is significant ($a_2\neq0$), the null hypothesis that there is a unit root has to be tested again, now using the normal distribution critical values (Enders 1995, 256). If the trend was not significant and we did not reject the unit root when testing with the intercept only, we checked whether the intercept itself is significant (test $a_0=0$ given $\gamma=0$). If the intercept was not significant, we repeated the ADF without trend and without intercept, using as many lags of ΔY_{1t} as needed. When needed, we reported all values of ADF, and indicated when the trend and the intercept were significant, writing the test statistics for the trend and the intercept below the values of the ADF tests enclosed in square brackets (see Tables A-1 and A-2).

Autoregressive Model. We fit a univariate autoregressive model for each series. This information complements the determination of the number of lags to include in the VEC model and provides us extra information about the existence of unit roots.

Confidence Intervals for the Impulse Response Functions obtained from VEC models.

We implemented a routine based on source code available from RATS for stationary VAR models. Scholars wishing to use this procedure should:

a) Estimate the VEC model using CATS. Obtain the impulse-response function for the original variables in levels using the coefficients estimated in CATS. This required modifying the source code of CATS to define as a global variable the array that stores the short-term matrices Γ_j, j=1,2,…,p-1. Then, obtain the coefficients of the VAR in levels from the short-term matrices and the error correction term coefficient, as follows:

$A_1 = \alpha\beta' + I + \Gamma_1, \quad A_j = \Gamma_j - \Gamma_{j-1}, j = 2,\ldots,p-1, \quad A_p = -\Gamma_{p-1}.$

From these A_j matrices, we can obtain the impulse-response matrices, but we do not have a measure of uncertainty of those estimates, so we need to estimate a VAR in differences with extra terms to obtain the confidence intervals for the impulse-response functions.

b) Take the values of the estimated cointegration vectors (β) as known (Hansen and Juselius 1995, 49) and estimate a VAR in differences with the $\beta'Y_{t-1}$ terms as extra exogenous variables. [17]

c) Obtain the impulse-response function from that model. This impulse-response function is identical to the impulse-response function we get directly from the CATS estimates as indicated in (a). In fact, our program obtains both to make sure they are identical.

d) Generate replicates of the estimated coefficients of the VAR in differences plus extra variables, using the variance matrix of those estimates in the same form as the program MONTEVAR.prg provided with RATS. This is a Monte Carlo method, sampling from the posterior distribution of the variance of the estimates (Wishart) and given that matrix, sampling from the posterior distribution of the estimates (Normal).[18]

e) Obtain the impulse-response function for each replicate, store all these functions, and obtain the 5^{th} and 95^{th} percentiles, giving an empirical 90% posterior-probability interval at each lead time. We obtain responses to orthogonal shocks using the Choleski decomposition of the variance-covariance of the error term.

f) Graph the impulse-response function together with the percentiles joining adjacent lead times to provide the boundaries.

The code used in this analysis is available from the authors. It requires the use of the CATS/RATS software.

[17] Enders (1996, 198) suggests this approach in an exercise.

[18] We thank Tom Maycock, from Estima, for his suggestions on how to implement this new program.

Table 1.

Partial Results from the Impulse Response Function of Vector Error Correction Models

Quarterly Data 1981:1 to 1998:3, T = 71 observations.

All models include eight variables: LRGDP, INVEST, FDIin, FDIout, TRADEtotal, TAX, CRATIO and PRES. All models include p = 3 lags in levels (2 lags in error correction form), and r = 6 cointegrating equations.

Impulse Response Function Results for Political Action Committees

(full sets of results available from the authors)

Variable used as TAX	TAX/Assets	TAX/GDP	TAX/EBIT	TAX/EBITDA
Resp. of PAC to GDP	ns	ns	ns	- 1
Resp. of PAC to INVEST	ns	ns	ns	ns
Resp. of PAC to FDI in	ns	ns	- 1	ns
Resp. of PAC to FDI out	+ 0:2	+ 0:1	+ 0, 1, 8	+ 0:1
Resp. of PAC to TRADE	+ 0:1	+ 0	ns	+ 0
Resp. of PAC to TAX	ns	ns	+ 1:6	+ 1
Resp. of PAC to President	+ 2:4	+ 2	+ 2	+ 2:4
Resp. of GDP to PAC	- 3:8	- 3:8	- 3:8	- 3:8
Resp. of INVEST to PAC	- 3:8	- 4:8	- 4:7	- 4:6
Resp. of FDI in to PAC	+ 1	+ 1	+ 1, - 4, 6:8	+ 1, - 6:8
Resp. of FDI out to PAC	- 4	+ 1, - 4	- 4	- 4
Resp. of TRADE to PAC	ns	ns	Ns	ns
Resp. of TAX to PAC	- 2:7	- 3	Ns	ns
Resp. of President to PAC	ns	ns	Ns	ns

The notation 0:3 means there is a significant response in lags 0, 1, 2 and 3. The sign indicates the direction of the effect, the numbers indicate the lags where the response is significant (i.e. where the boundaries do not contain the value zero). The boundaries are given by the 5^{th} and the 95^{th} percentiles of the impulse-response functions obtained by Monte Carlo using 2000 draws. We only report responses up to lag 8. Non-significant responses are indicated with "ns".

Table A-1. Univariate Analysis, Quarterly Data
Augmented Dickey-Fuller Unit Root Tests (number of lags of ΔY included)
Sample 1954:1 to 1980:4, T=108 observations

Series	Intercept and trend [t-stat for trend]	Intercept only [t-stat for intercept]	None	Conclusion
TAX/GDP	-2.49 (8) [-2.20]	-1.20 (8) [0.94]	-1.85 (8)	I(1)
TAX/Assets	-1.91 (8) [-1.88]	0.50 (8) [0.22]	-1.70 (8)	I(1)
TAX/EBIT	-2.22 (8) [-0.69]	-2.70 (8) [2.52]	-1.83 (8)	I(1)
TAX/EBITDA	-3.31** (8) [-2.80*]			I(0) with - trend

Table A-2. Univariate Analysis, Quarterly Data
Augmented Dickey-Fuller Unit Root Tests (number of lags of ΔY included)
Sample 1981:1 to 1998:3, T=71 observations

Series	Intercept and trend [t-stat for trend]	Intercept only [t-stat for intercept]	None	Conclusion
LRGDP	-2.60 (2) [2.63]	.18 (1) [-0.09]	3.79 (1)	I(1)
Invest = PDE/Assets	-1.75 (5) [2.28]	.03 (5) [0.15]	1.26 (5)	I(1)
TAX/GDP	-4.02* (7) [2.30*]			I(0) with + trend
TAX/Assets	-3.75* (7) [3.22*]			I(0)with + trend
TAX/EBIT	-4.98**(7) [-0.03]			I(0)
TAX/EBITDA	-5.04**(7) [2.15*]			I(0) with + trend
FDI in/GDP	-2.48 (6) [1.26]	-2.16 (6) [2.18]	-.37 (6)	I(1)
FDI out/GDP	-8.26** (0) [6.34**]			I(0) with + trend
TRADE total /GDP	-3.03 (2) [3.02*]			I(0) with + trend
Cratio of PAC (Corporate/Total)	-2.79 (2) [-1.62]	-2.57 (2) [2.52]	-1.61 (2)	I(1)
PRES = President	-4.67**(0) [4.73**]			Dummy Variable
CONGRESS (%)	-1.46 (0) [-.65]	-1.35 (0) [-.27]	-1.58 (0)	I(1)
President & Congress (%)	-5.80** (0) [5.13**]			I(0) with + trend

* .01 < p-value < .05, ** p-value < .01, using MacKinnon critical values for rejection of the null hypothesis of a unit root, and Table 4.1 in Enders, 1995, 223 for testing the trend and the intercept.

Note: When the trend is significant like in the case of TAX/EBITDA and TRADE, the test statistic for the unit root has to be compared against the critical values of the normal distribution, and in those cases (-3.31 and -3.03) there is evidence to reject the null hypothesis of a unit root (Enders p.256).

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