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Coordinating Economist:
Randall W. Eberts

Editors: Tess Ferg
Robin Ratliff
Design: Michael Galka
Typography: Liz Hanna

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Bank Capital Requirements and Leverage: A Review of the Literature

by William P. Osterberg

William P. Osterberg is an economist at the Federal Reserve Bank of Cleveland.
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Introduction

Recognition of the extensive losses inflicted on the Federal Savings and Loan Insurance Corporation (and thus indirectly on the taxpayers) by the thrift industry crisis has led to increased scrutiny of the safety and soundness of commercial banks and other financial institutions. It has become obvious that some of the factors responsible for excessive risk-taking by savings and loans may also be relevant to commercial banks. In particular, the current system of fixed-rate deposit insurance and supervision and regulation interacts in complex ways with the market forces that may ordinarily discipline banks. An understanding of these interactions is crucial to financial institution reform.

This article reviews the literature relevant to assessing one proposed regulatory reform — increased capital requirements for banks. The arguments for higher capital requirements rely primarily on the premise that they will strengthen market discipline, and secondarily on the desire to provide a greater cushion for the deposit insurance agency. In theory, increased capital requirements can at least partially compensate for the weakening of market discipline that may result from the continued presence of fixed-rate deposit insurance. However, the magnitude of the impact of past changes in capital requirements on banks' capital decisions is unclear, mainly because market forces also affect such decisions.

Any analysis of the impact of capital requirements must take into account the current system of fixed-rate deposit insurance. The presence of non-risk-related insurance premiums and government guarantees influences the capital decision by blunting the effect that increased capital-asset ratios would otherwise have on banks' cost of funds. The current system complicates identifying the impact of changes in required capital-asset ratios, because the subsidy itself may be influenced by the ratios and other factors.

Fixed-rate insurance is also widely viewed as subsidizing risk-taking, thus providing a rationale for capital regulation. In the absence of government guarantees, shareholders would need higher levels of capital as a buffer against losses in order to avoid risk-related increases in their cost of funds. These guarantees thus lead to a substitution of deposits for equity, thereby lowering capital ratios.

Although this distortion has led to reform proposals that emphasize reductions in the scope of government guarantees, proposals to increase capital requirements continue to emerge despite the introduction of risk-based requirements. This has occurred in part because changes in capital requirements are seen as relatively easy to implement. In addition, as noted above, capital requirements may induce shareholders to evaluate risk more carefully and to submit to the market's evaluation when they attempt to raise capital.

On the surface, capital requirements seem to be effective, because almost all banks increase their capital-asset ratios (book value) after the requirements are increased. However, other factors may influence the ratios, especially if they are calculated in terms of book value. For example, suppose that the regulatory standards were increased in response to a general market perception that capital is inadequate. In this case, the subsequent adjustment may be partly due to banks' desire to avoid an increase in the risk premium in their cost of funds. Clearly, in order to disentangle such influences, investigators must have a model of the factors that determine bank capital-asset ratios.

The remainder of this paper is organized as follows. Section I reviews the theoretical literature on the determination of banks' capital structure. Section II covers a closely related topic, the impact of capital requirements on portfolio risk. Section III presents a model of a bank's choice of capital structure. Section IV analyzes the model's implications for the impact of market and regulatory forces on bank leverage.[1] Section V reviews the empirical findings on the effects of capital requirements, contrasting various results in terms of the implications presented in section IV. Section VI presents suggestions for future research and concludes.

I. Optimal Capital Structure Theory for Financial Institutions

I first discuss the theoretical findings relevant to nonfinancial corporations, since to some extent these may extend to banks, and then review the limited number of analyses of how and why the capital structure decisions of banks may differ from those of nonfinancial institutions.[2] I then focus on specific analyses of banks' capital structure, most of which assign a prominent role to deposit insurance.

The literature analyzing the capital structure decisions of nonfinancial corporations is so broad as to defy easy description.[3] However, one of the strongest conclusions to emerge from the empirical studies is that optimal capital structure is influenced by the tax code, possibly in combination with leverage-related costs. For example, the ability of corporations to deduct interest on debt may encourage an increase in leverage. On the other hand, higher levels of nondebt tax shields, such as depreciation and tax credits, may reduce optimal leverage by increasing the probability that not all interest expenses will be deductible. Taxes on personal equity and interest income may also decrease the net tax advantage of debt and optimal leverage.

Leverage-related costs include the expected costs of bankruptcy, and agency costs associated with conflicts among creditors, stockholders, and managers. The direct costs of bankruptcy are minimal (mainly involving administrative and legal fees), but agency costs, which include any decrease in firm value associated with contractual arrangements to protect one party from actions taken by another party with conflicting interests, can be significant. Bond covenants that restrict cash-flow usage may impose agency costs. However, such covenants may be a part of optimal contracts reconciling bondholders with stockholders.

The theory of optimal financial structures for financial intermediaries differs somewhat from the theory for nonfinancial firms. First, in analyzing capital structure for either financial or nonfinancial firms, it is convenient to assume that operating and financing decisions can be separated. This assumption is harder to defend for financial intermediaries. The existence of complete markets, which makes separation

■ 1 Leverage is often defined as the ratio between debt and equity, measured in book or market values. In the model presented here, the bank chooses the level of promised payments to depositors, given an exogenous asset portfolio. This is equivalent to choosing the debt-to-equity ratio directly. In Osterberg and Thomson's (1990) empirical study, the measure of leverage is the ratio between the book value of debt and the total of the book value of debt and the market value of equity. This is close to another often-analyzed measure of leverage, the debt-to-asset ratio.

■ 2 In this section, we review the theoretical analyses relevant to understanding the impact of changes in bank capital requirements. Because few analyses of bank capital structure are available (relative to the number that deal with the capital structure of nonfinancial institutions), it is not useful to attempt to categorize various approaches. In addition, the dissimilarities in approach prevent the development of a general model to which all others specialize.

■ 3 Among several useful surveys is one by Harris and Raviv (1990), who categorize the forces that may influence capital structure into desires to 1) ameliorate conflicts of interest, 2) convey private information to markets, 3) influence product or input markets, and 4) affect corporate control contests. The authors exclude tax-driven theories that they admit are of great empirical importance. Although few analyses of the capital structure of banks consider these four forces, several take into account taxes and other considerations discussed here.

more likely, makes it difficult to explain the existence of intermediaries: If markets were complete, lenders and borrowers could transact without them. In addition, deposits seem to play a role in both the real and financial decisions of banks, because deposits are not only an input into banks' production, but a component of debt in their capital structure. Another reason that analyses of banks' capital structure differ is that regulatory forces aimed directly at capital structure (for example, capital-asset ratios) must be considered.

Although most studies of the impact of capital requirements on banks do not view these institutions as fundamentally different from nonfinancial entities, many others have examined the role of informational asymmetries and contracts in explaining the existence of intermediaries. Early examples are Boyd and Prescott (1986) and Diamond (1984). Sealey (1985) analyzes a model of incomplete markets and intermediaries, showing the conditions under which shareholder unanimity holds and under which unanimity implies separation. Sealey (1983) examines a model of incomplete markets in which economies of scale in the provision of deposit services influence bank leverage. Chen, Doherty, and Park (1988) utilize an option-pricing framework to analyze the capital structure decisions of depository financial intermediaries in the presence of deposit insurance, reserve requirements, liquidity effects, and taxation. They conclude that no clear separation exists between operating and financial decisions, and that this finding even applies to analyses of the impact of taxation on leverage decisions.

As noted in Santomero (1984), most studies of bank capital structure assume that real and financial decisions can be separated, and try to explain leverage choice conditional on a given portfolio of assets. One example is Orgler and Taggart (1983), who show how personal and corporate taxes, reserve requirements, and economies of scale influence intermediaries' optimal leverage. Applications of the option-pricing framework also assume that portfolio composition is held constant. Pyle (1986) shows that the use of book values in capital regulation is inappropriate when combined with closure rules that deviate from an economic solvency condition.

The conclusions of theoretical analyses of the impact of capital requirements are closely related to the treatment of deposit insurance and government guarantees. If deposit insurance is underpriced and unresponsive to risk, then stockholders are being subsidized by the insurer, and the size of the subsidy is a function of portfolio risk and leverage. This subsidy has a direct impact on banks' responses to changes in capital requirements. Buser, Chen, and Kane (1981) examine how the combination of capital regulation and flat-rate deposit insurance jointly influences bank leverage. They note that because capital regulation encompasses more than just numerical standards for capital-asset ratios, such regulation can be seen as imposing an implicit risk-related insurance premium that discourages banks from exploiting the subsidy implied by flat-rate deposit insurance.

II. The Impact of Capital Requirements on Portfolio Composition

Although most studies of bank capital structure assume a given portfolio of assets, several authors have examined the impact of capital requirements on portfolio risk, assuming that leverage is at the regulatory maximum.[4] An overall assessment of the impact of capital requirements on bank capital structure would have to allow for possible feedback from variation in portfolio changes.

Koehn and Santomero (1980) conclude that increased numerical capital requirements lead banks that are risk-averse expected utility maximizers to reshuffle their portfolios so as to increase the probability of bankruptcy. Lam and Chen (1985) and Kim and Santomero (1988) use similar approaches. Keeley and Furlong (1987), who employ a value-maximization framework, point out that Koehn and Santomero ignore the impact of changes in leverage and portfolio risk on the deposit insurance subsidy. Osterberg and Thomson (1988) show how the impact of capital requirements on portfolio shares is altered by allowing the cost of funds to be influenced by leverage.

4 Flannery (1989) shows why insured banks may have a preference for safe individual loans but still prefer risky overall portfolios. Capital adequacy standards and loan examination procedures are key elements of his analysis. Lucas and McDonald (1987) study the impact of capital regulation on bank portfolio choice when banks have private information about loan quality.

EQUATIONS (1) AND (2)

$$(1)\quad \tilde{Y}_s = (\tilde{X} - \hat{Y})(1 - t_c) + \varphi \qquad \text{if } \tilde{X} \geq \hat{Y} + \frac{\delta - \varphi}{1 - t_c}$$

$$= (1 - \lambda)[(\tilde{X} - \hat{Y})(1 - t_c) + \varphi] \qquad \text{if } \hat{Y} + \frac{\varphi}{t_c} \leq \tilde{X} < \hat{Y} + \frac{\delta - \varphi}{1 - t_c}$$

$$= (1 - \lambda)(\tilde{X} - \hat{Y}) \qquad \text{if } \hat{Y} \leq \tilde{X} < \hat{Y} + \frac{\varphi}{t_c}$$

$$= 0 \qquad \text{if } \tilde{X} < \hat{Y}$$

$$(2)\quad \hat{Y}_b = \hat{Y} \qquad \text{if } \hat{Y} \leq \tilde{X}$$

$$= \tilde{X}(1 - k) \qquad \text{if } 0 < \tilde{X} < \hat{Y}$$

$$= 0 \qquad \text{otherwise,}$$

where

$\tilde{X}$ = end-of-period value of bank assets,

$\tilde{Y}_s, \tilde{Y}_b$ = gross end-of-period cash flows accruing to bank stockholders and depositors, respectively,

$\hat{Y}$ = total end-of-period promised payment to depositors,

φ = total end-of-period after-tax value of nondebt tax shields when fully utilized,

λ = regulatory penalty,

δ = capital requirement,

k = cost of financial distress to depositors, and

d = proportion between δ and $\tilde{Y}$ (a capital requirement proxy).

III. A Model of Market and Regulatory Influence on Bank Capital Structure

To aid in this review, I present a model in which market and regulatory influences on banks' capital structure are intertwined. The model also provides a limited synthesis of the theoretical literature. However, influences that could explain the existence of intermediaries, such as the presence of incomplete markets, are not incorporated. The only factor included that distinguishes banks is capital regulation. In addition, the model maintains the separation of real and financial decisions by holding constant the bank's asset portfolio and return variance. Although I initially assume that there is no deposit insurance, such insurance is easily introduced (see Osterberg and Thomson [1990] and the following discussion).[5]

Equations (1) and (2) describe the uncertain outcomes facing stockholders and depositors. I view the bank as attempting to maximize the total of the values of equity and deposits, each of which depends on the uncertain outcomes and their associated probabilities.

Pre-tax returns to stockholders depend on the uncertain end-of-period value of bank assets, $\tilde{X}$. The first line of equation (1) indicates the return when income (asset values) is high enough that the capital guideline is not violated. I assume in this case that all nondebt tax shields can be utilized ($\tilde{X} > \hat{Y} + \varphi/t_c$); however, the results are not significantly affected by this assumption. In the second case, when income is high enough to use all the shields but the capital requirement is not met ($\tilde{X}[1 - t_c] + \varphi - \hat{Y} < \delta$), regulators impose a tax of λ on stockholder returns. In the third case, income is positive but insufficient to utilize nondebt tax shields, and the guidelines are not met.

5 This model is a variant of the one developed by Bradley, Jarrell, and Kim (1984), hereafter referred to as BJK. Detailed assumptions underlying the model are given in appendix 1.

Equation (2) indicates the end-of-period pretax flows to depositors. A crucial distinction between stockholders and depositors is readily apparent: Depositors only receive $\hat{Y}$, even if income greatly exceeds promised payments. On the other hand, if income is positive but insufficient to meet promised payments, the bank is in financial distress and incurs real costs that reduce the return to depositors by the fraction k.

The bank is assumed to know 1) the relevant tax rates, 2) the amount of nondebt tax shields, 3) the required capital-asset ratio, d, 4) the regulatory response, embodied in λ, 5) the costs of financial distress, 6) the average income, $\tilde{X}$, and 7) the standard deviation of income, φ. The bank chooses $\hat{Y}$ to maximize the market value of its debt plus equity (see appendix 1).

Equation (3) is the derivative of V with respect to $\hat{Y}$: $V_{\hat{Y}}(\partial V / \partial \hat{Y})$.

$$
\begin{aligned}
(3) \quad V_{\hat{Y}} = {} & \frac{1 - t_{pb}}{r_0} [1 - F(\hat{Y}) - k\hat{Y} f(\hat{Y})] \\
& + \frac{1 - t_{ps}}{r_0} \Big[-(1 - t_c)[1 - F(\frac{\varphi}{t_c} + \hat{Y})] \\
& - [F(\hat{Y} + \frac{\varphi}{t_c}) - F(\hat{Y})] - \lambda \{ [F(\hat{Y} + \frac{\varphi}{t_c}) \\
& - F(\hat{Y})] + (1 - t_c)[F(\hat{Y} + \frac{\delta - \varphi}{1 - t_c}) \\
& - F(\hat{Y} + \frac{\varphi}{t_c})] + [\delta + \frac{d(\delta - \varphi)}{1 - t_c}] \\
& \cdot f(\hat{Y} + \frac{\delta - \varphi}{1 - t_c}) \} \Big],
\end{aligned}
$$

where $F(\cdot)$ is the cumulative probability density function of $\tilde{X}$. If banks in fact choose $\hat{Y}$ so as to satisfy equation (3), then this expression indicates how both market and regulatory forces influence bank leverage.

If $\lambda = 0$, the model's implications are consistent with theories of optimal capital structure in which the assumed tax advantage of debt balances the expected cost of bankruptcy (see BJK). These implications are as follows. First, an increase in t_{ps} raises optimal leverage by increasing the cost of equity. Analogous reasoning implies that an increase in t_{pb} reduces optimal leverage. Second, an increase in t_c raises optimal leverage by increasing the tax advantage of debt. For this reason, an increase in φ reduces optimal leverage by increasing the probability that not all interest expenses will be deductible. Third, an increase in k reduces optimal leverage by increasing the expected cost of a bank's inability to make all promised payments.

The model is also consistent with theoretical approaches that assign deposit insurance a role in distorting market discipline. The effect of fixed-rate deposit insurance on optimal leverage can be seen by comparing equation (3) with $V_{\hat{Y}}$ under full insurance (see Osterberg and Thomson [1990]). Optimal leverage is higher with fixed-rate deposit insurance by the amount $(1 - t_{pb})[F(\hat{Y}) + k\hat{Y}f(\hat{Y})]/r_0$. Fixed-rate deposit insurance increases the optimal $\hat{Y}$ by insuring that depositors are always paid in full and by shifting the cost of financial distress from depositors to the Federal Deposit Insurance Corporation. In the context of the model presented above, these two influences are equivalent to assuming that $F(\hat{Y}) = 0$ and $k = 0$. Although this model does not allow the higher leverage to influence bank riskiness, the increase in leverage induced by deposit insurance provides a rationale for capital regulation.

The impact of capital regulation on leverage can be seen by examining the last term in equation (3), $\lambda(\cdot)$. The first two components of $\lambda(\cdot)$ comprise the expected after-tax regulatory penalty resulting from issuing the last dollar of deposits. As equations (1) and (2) demonstrate, the possibility of a regulatory penalty affects the return to equity, which one would expect to be reflected in the rate of return demanded by stockholders and thus in the bank's leverage decision. In fact, the last component of $\lambda(\cdot)$ is the increase in the cost of equity capital that results from issuing one more dollar of deposits, $[\delta + d(\delta - \varphi)/(1 - t_c)] f[\hat{Y} + (\delta - \varphi)/(1 - t_c)]$ Because all of the components are positive, the possibility of a regulatory penalty reduces a bank's optimal leverage.

IV. The Impact of Regulatory and Market Forces on Optimal Leverage

Although equation (3) clearly shows that both market forces and regulatory variables influence leverage with signs consistent with theory, it is more important for our purposes to note that this expression also implies that the impact of an increase in λ (the regulatory penalty) on leverage depends on market forces entering $\lambda(\cdot)$. Empirical studies of capital requirements vary in their treatment of the influence of such market forces (φ, k, t_c, t_{ps}, t_{pb}, and σ, where σ is the standard deviation of $\tilde{X}$) on $\hat{Y}$.

To show how market influences affect the impact of capital regulation on bank leverage, one can differentiate the optimality condition (equation [3]) with respect to the regulatory variables. The derivatives with respect to the market-force variables are indicated in appendix 2. Further details can be found in Osterberg and Thomson (1990).

Equation (4) gives the impact of a change in d on optimal leverage. The ratio d is closely related to a required capital-asset ratio, because it is the minimum level of the end-of-period equity value and because $\delta = \hat{Y}d$.

$$(4) \quad V_{\hat{Y}d} = -\frac{\lambda(1-t_{ps})}{r_0(1-t_c)} f(\hat{Y} + \frac{\delta-\varphi}{1-t_c}) \{2\delta - \varphi - [\hat{Y}\delta + \frac{\delta(\delta-\varphi)}{1-t_c}] (\hat{Y} + \frac{\delta-\varphi}{1-t_c} - \bar{X})/\sigma^2\} \gtrless 0$$

The impact of d on leverage clearly depends on market forces, implying that such forces influence leverage even if a bank fails to meet the guidelines. As discussed below, some studies imply that such banks are influenced only by regulation, while banks meeting the guidelines are influenced only by market forces. No such dichotomy emerges here.

Equation (4) implies that $V_{\hat{Y}d}$ is negative whenever $\bar{X} \geq \hat{Y} + (\delta - \varphi)/(1 - t_c)$; that is, an increase in d reduces leverage when the bank expects to meet the capital requirements. However, if a bank does not expect to meet the requirements, an increase in d may induce it to increase leverage and thus move even further below the guidelines.

Equation (5) shows that an increase in the regulatory penalty, λ, reduces bank leverage. Here, as in the response of leverage to d, the impact of capital regulation depends on market factors. Equation (6) shows that an increase in the costs of financial distress, k, also reduces optimal leverage. Although k is referred to above as a market factor, the cost of financial distress can be influenced by regulatory policies pertaining to bank closure.

$$(5) \quad V_{\hat{Y}\lambda} = -\frac{1-t_{ps}}{r_0} \Big[(1-t_c) [F(\hat{Y} + \frac{\delta-\varphi}{1-t_c}) - F(\hat{Y} + \frac{\varphi}{t_c})] + (\delta + \frac{d(\delta-\varphi)}{1-t_c}) f(\hat{Y} + \frac{\delta-\varphi}{1-t_c})\Big] < 0$$

$$(6) \quad V_{\hat{Y}k} = -\frac{1-t_{pb}}{r_0} [\hat{y} f(\hat{Y})] < 0$$

V. Evidence on the Impact of Capital Requirements on Bank Leverage

Separating market forces from regulatory forces has been a major difficulty in ascertaining the effectiveness of capital guidelines. Dietrich and James (1983) criticize earlier studies by Peltzman (1970) and Mingo (1975) for ignoring deposit-rate ceilings in their analyses of the impact of capital requirements. Under such ceilings, banks can influence risk-adjusted returns on bank debt by augmenting capital. However, only under less-than-full deposit insurance would more capital benefit stockholders, by inducing uninsured depositors to accept lower interest rates. Dietrich and James conclude that the guidelines have no effect on bank capital changes.

Although the model presented here does not directly consider the possibility of interest-rate ceilings, capital levels influence the returns to stockholders and thus the rate of return required on equity. The latter can be calculated as the ratio between $E(\hat{Y}_s)$, the expected returns to stockholders, and S, the market value of equity (see appendix 1 and BJK). Equation (1) indicates that returns to stockholders are influenced by several market forces that must be controlled for in any analysis of the impact of capital requirements.

Marcus (1983), Wall and Peterson (1987), and Keeley (1988a, 1988b) examine bank holding companies rather than independent banks. Wall and Peterson apply a switching regression technique to movements of equity values in an attempt to distinguish a regime in which capital ratios exceed the requirements (and are thus influenced by market forces) from a regime in which ratios are at the regulatory limit. They conclude that most banks are influenced by regulation.

The model presented here implies that 1) banks may respond to market forces even if the guidelines are not being met and 2) regulatory forces may influence leverage even if the bank exceeds the guidelines. In addition, equation (4) indicates that banks below the guidelines may actually respond to stiffer requirements perversely.

Keeley (1988a) examines the response of bank holding companies to the increased capital requirements of the 1980s. Although capital-deficient banks increased their book-value ratios more than capital-sufficient banks did, market ratios increased for both classes. However, regulatory subsidies or taxes can influence the response of market-value ratios to increased capital guidelines, because the value of the subsidy may vary with leverage or risk. Keeley (1988b) claims that increased competition erodes the value of bank charters and thus raises incentives to increase leverage or to reduce capital ratios.

Marcus (1983) utilizes a time series cross-sectional approach, measuring regulatory pressure to increase capital by the holding company's capital ratio relative to the average (in terms of book or market value). He finds that the incentive to decrease capital varies positively with the level and variability of interest rates, as well as with the tax disadvantage of equity finance. Regulation seems to have no effect. However, his regulatory measure does not incorporate risk.

In the model presented above, d is close to a statutory capital-asset ratio. However, analyzing banks' capital ratios relative to the average may be a more useful way to isolate the impact of capital regulation. There are at least two reasons for this. First, relatively few banks are below the statutory guidelines. Second, evidence suggests that capital regulation is based on a peer-group standard. In fact, a peer-group capital standard may be a useful proxy for the regulatory penalty variable, λ.

The relevance of taxes to the capital structure of banks is discussed in more detail by Wall and Peterson (1988) and Gelfand and Hanweck (1987). Wall and Peterson argue that taxes do not influence the capital structure of banks affiliated with holding companies, because the tax consequences of the parent issuing debt to buy subsidiary equity are similar to those ensuing when the bank itself issues debt. Gelfand and Hanweck examine the financial statements of 11,000 banks and find strong evidence for market influences (tax rates, risk, and municipal securities [munis] as proxies for nondebt tax shields) on leverage.

Osterberg and Thomson (1990) investigate the influence of capital regulation on bank holding company leverage empirically, drawing on the implications of the model presented above. The authors find that market forces influence leverage through three channels: a direct channel, a channel in which market forces interact with risk (σ), and a channel in which market forces interact with capital regulation. In addition, their analysis explicitly allows for the simultaneous determination of leverage and muni holdings. Although the latter may no longer be an important channel through which banks manage their tax liability, this may not have been the case during the period examined (1986-1987).[6] The interactive capital regulation measures, taken as a whole, are significant, as are the interactive risk measures. In addition, muni holdings appear to be significant determinants of leverage, as do market forces.

VI. Conclusions and Suggestions for Future Research

This article reviews the literature relevant to assessing the impact of increased bank capital requirements. Although researchers have suggested various proposals to correct the distorted incentives facing bankers, raising required capital ratios continues to emerge as a possible means of strengthening market discipline. However, previous studies have failed to clarify the impact of numerical guidelines on banks' capital-asset ratios.

The primary difficulty in discerning the influence of such guidelines lies in disentangling the impacts of regulatory and market forces. In order to illustrate the way in which these forces interact, I present a model of a bank's choice of leverage ratio where, in the absence of capital regulation, tax considerations and bankruptcy costs imply an interior solution. When capital regulation is introduced, it becomes clear that the impact of such regulation depends on market forces.

These results may provide useful insight for regulators. For example, the response of bank leverage to capital regulation may depend on

■ 6 Scholes, Wilson, and Wolfson (1990) present evidence that banks' muni holdings responded to changes in the tax code between 1983 and 1987, and that capital regulation seemed to influence banks' timing of capital loss realization. This seems to suggest that capital regulation and the tax code interact in a manner similar to that suggested in this paper.

the market factors considered in this paper, such as tax rates, nondebt tax shields, and muni holdings, not just on the capital position of the bank. This implies that evaluations of banks' leverage and capital-asset ratios should take into account market influences on the leverage decision.

The model may also explain previous empirical findings regarding the impact of capital requirements. Most studies do not control for many of the market influences on banks' capital decisions. The analysis presented here thus implies that theoretical examinations of bank capital structure may further improve our understanding of the influence of capital requirements. In this regard, it may be particularly useful to analyze capital requirements through models that incorporate informational asymmetries and market imperfections to explain the existence of financial intermediaries.

Appendix 1

Detailed Assumptions and Structure of the Model

The main assumptions of the model presented in the text are as follows:

1. Investors are risk-neutral.
2. The personal tax rates on returns from bank debt and bank equity are t_{pb} and t_{ps}, respectively.
3. Bank income is taxed at the corporate rate, t_c.
4. All taxes are levied on end-of-period wealth.
5. The firm's end-of-period tax liability can be reduced through nondebt tax shields, φ, such as investment tax credits and depreciation.
6. Unused tax credits cannot be transferred across time or across firms.
7. If banks cannot meet their end-of-period promised payments to depositors, $\hat{Y}$, costs of financial distress are incurred that reduce bank equity value by a factor of k.
8. The end-of-period capital requirement is $\delta = \hat{Y}d$.
9. If $\tilde{X} - \hat{Y} < (\delta - \varphi)/1 - t_c)$, a regulatory penalty reduces stockholders' returns by a constant fraction λ ($\tilde{X}$ is the end-of-period value of assets).
10. All bank liabilities are *uninsured* deposits.
11. The capital constraint, δ, is not binding unless $\tilde{X}$ is such that the tax shields are being fully utilized.

Assumption 10 allows us to separate the effects of capital requirements from the effects of deposit insurance. Thomson (1987) shows that this is equivalent to assuming 100 percent deposit insurance if the insurance is fairly priced. The case in which all liabilities are covered by fixed-rate, zero-premium deposit insurance is analyzed in appendix A of Osterberg and Thomson (1990). Assumption 11 is made for convenience only; my results are not materially affected by the alternative assumption that δ is binding for values of $\tilde{X}$ where $\varphi > (\tilde{X} - \hat{Y})\, t_c$.

Under the assumption of risk neutrality, and given the uncertain outcomes detailed in the text, the after-tax market value of the banking firm is the sum of the market values of deposits and equity:

$$
\text{(1A)} \quad V = \frac{1}{r_0}\Big[\int_{\hat{Y}+\frac{\varphi}{t}}^{\infty} \{ (1-t_{ps}) [(\tilde{X}-\hat{Y})(1-t_c) + \varphi] + (1-t_{pb})\hat{Y} \} f(\tilde{X})\, d\tilde{X}
- \int_{\hat{Y}+\frac{\varphi}{t}}^{\hat{Y}+\frac{\delta-\varphi}{1-t}} \lambda (1-t_{ps}) [(\tilde{X}-\hat{Y})(1-t_c) + \varphi] f(\tilde{X})\, d\tilde{X}
+ \int_{\hat{Y}}^{\hat{Y}+\frac{\varphi}{t}} [(1-t_{ps})(1-\lambda)(\tilde{X}-\hat{Y}) + (1-t_{pb})\hat{Y}] f(\tilde{X})\, d\tilde{X}
+ \int_{0}^{\hat{Y}} (1-t_{pb})(1-k)\tilde{X} f(\tilde{X})\, d\tilde{X} \Big],
$$

where $f(\tilde{X})$ = the probability density of $\tilde{X}$, and r_0 = one plus the rate of return on a risk-free tax-exempt bond.

The four integrals in equation (1A) are, respectively, 1) the expected value of the bank over the range of $\tilde{X}$ where the bank fully utilizes its nondebt tax shields, 2) the expected value of the regulatory tax over the range of $\tilde{X}$ where the bank fully utilizes its nondebt tax shields but fails to meet its capital guideline, 3) the expected value of the bank over the range of $\tilde{X}$ where nondebt tax shields are no longer fully utilized, and 4) the expected value of the bank when $\tilde{X}$ is not large enough to meet promised payments to the depositors and k percent of the firm value is lost to financial distress.

Appendix 2

The Impact of Market Forces on Optimal Bank Leverage

The effect of an increase in nondebt tax shields, φ, on optimal leverage is indicated by equation (2A).

$$\text{(2A)} \quad V_{\hat{Y}\varphi} = -\frac{1-t_{ps}}{r_0}\Big[(1+\lambda)f(\hat{Y}+\frac{\varphi}{t_c}) + \lambda f(\hat{Y}+\frac{\delta-\varphi}{1-t_c})\{1+\frac{d}{1-t_c} - [\delta+\frac{d(\delta-\varphi)}{1-t_c}] \cdot (\hat{Y}+\frac{\delta-\varphi}{1-t_c}-\bar{X})/\sigma^2\}\Big] \gtrless 0$$

If there were no regulatory penalty ($\lambda = 0$), I would obtain the same results as BJK; that is, a higher level of nondebt tax shields would reduce leverage ($V_{\hat{Y}\varphi} < 0$). Here, however, leverage increases if $\bar{X} \geq \hat{Y} + (\delta-\varphi)/(1-t_c)$. This possibility is created by the combination of the capital requirement being based on the after-tax value of equity, which includes the value of the shields, and the fact that the capital requirement is binding when the tax shields are being fully utilized. For high-enough values of $\bar{X}$, an additional dollar of tax shields reduces the probability that the bank will violate the capital constraint and incur the regulatory penalty.

The effects of changes in the various tax rates on the optimal level of debt are shown in equations (3A), (4A), and (5A). In equation (3A), the response of bank leverage to an increase in the marginal corporate tax rate is positive when $\bar{X} \geq \hat{Y} + (\delta-\varphi)/(1-t_c)$. In other words, if expected end-of-period income is large enough to meet the capital requirements, then an increase in t_c reduces the optimal level of debt. The ambiguous sign for equation (3A) when $\bar{X} \geq \hat{Y} + (\delta-\varphi)/(1-t_c)$ arises because the capital constraint is assumed to be binding when the bank's net tax bill is positive. There are two offsetting effects. First, an increase in t_c raises the value of the interest deduction on debt, which induces the bank to issue more deposits. This is the familiar effect discussed in the finance literature on optimal capital structure for nonfinancial entities. The second effect is a reduction in the after-tax value of equity and an associated increase in the probability that the bank will violate the capital constraint and reduce leverage.

$$\text{(3A)} \quad V_{\hat{Y}t} = \frac{1-t_{ps}}{r_0}\Big\{1-F(\hat{Y}+\frac{\varphi}{t_c}) + \lambda[F(\hat{Y}+\frac{\delta-\varphi}{1-t_c}) - F(\hat{Y}+\frac{\varphi}{t_c})] + \frac{\varphi(1-\lambda)}{t_c}f(\hat{Y}+\frac{\varphi}{t_c}) + \Big[(1+\frac{d}{1-t_c})[\frac{\lambda(\delta-\varphi)}{1-t_c}] - \frac{\lambda}{\sigma^2} \cdot [\delta+\frac{d(\delta-\varphi)}{1-t_c}](\hat{Y}+\frac{\delta-\varphi}{1-t_c}-\bar{X})\Big] \cdot f(\hat{Y}+\frac{\delta-\varphi}{1-t_c})\Big\} \gtrless 0$$

If there were no costs of financial distress ($k=0$), equation (4A) would be unambiguously negative at the optimal level of debt. In addition, if all of the bank's deposits were insured, $V_{\hat{Y}t_{pb}}$ would be clearly negative. However, more generally, equation (4A) is negative when the probability that $\hat{Y}$ is less than $\tilde{X}$ exceeds the marginal expected leverage-related costs. This result is similar to the findings of BJK. Note that we have assumed that the costs of financial distress facing the depositors (k) are distinct from the regulatory penalty. As in BJK, $V_{\hat{Y}t_{ps}}$ is unambiguously positive. However, here the response depends on the regulatory penalty, λ.

$$\text{(4A)} \quad V_{\hat{Y}t_{pb}} = -\frac{1}{r_0}[1-F(\hat{Y}) - k\hat{Y}f(\hat{Y})] \gtrless 0$$

$$(5A)\quad V_{\hat{Y}t_{ps}} = -\frac{1}{r_0}\Big\{(1-t_c)\,[\,1-F(\frac{\varphi}{t_c}+\hat{Y})\,]$$

$$+\,[\,F(\hat{Y}+\frac{\varphi}{t_c}) - F(\hat{Y})\,]$$

$$+\,\lambda\Big[\,[\,F(\hat{Y}+\frac{\varphi}{t_c}) - F(\hat{Y})\,]$$

$$+\,(1-t_c)\,[F(\hat{Y}+\frac{\delta-\varphi}{1-t_c}) - F(\hat{Y}+\frac{\varphi}{t_c})\,]$$

$$+\,(\delta+\frac{d(\delta-\varphi)}{1-t_c})\,f(\hat{Y}+\frac{\delta-\varphi}{1-t_c})\,\Big]\Big\} > 0$$

Finally, the optimal level of deposits is a function of the variability of $\tilde{X}$. Equation (6A) shows that an increase in σ has an ambiguous effect on optimal leverage. The sign on $V_{\hat{Y}\sigma}$ depends on the proximity of $\hat{Y}$, $\hat{Y} + (\delta - \varphi)/(1-t_c)$, and $\hat{Y} + \varphi/t_c$ to the mean of $\tilde{X}$, as well as on the magnitudes of k, φ, d, and λ. BJK find that, even without a regulatory penalty, the impact of an increase in σ on $\hat{Y}$ is ambiguous.

$$(6A)\quad V_{\hat{Y}\sigma} = \frac{1-t_{pb}}{r_0\sigma}$$

$$\cdot\,\{\,(\hat{Y}-\bar{X}) - k\hat{Y}[\,(\frac{\hat{Y}-\bar{X}}{\sigma})^2-1\,]\,\}\,f(\hat{Y})$$

$$+\,\frac{1-t_{ps}}{r_0\sigma}\Big\{\,[\,(1-\lambda)\,t_c - 2\lambda\,]$$

$$\cdot\,(\frac{\varphi}{t_c}+\hat{Y}-\bar{X})$$

$$\cdot\,f(\hat{Y}+\frac{\varphi}{t_c}) - (1-\lambda)\,f(\hat{Y})\,(\hat{Y}-\bar{X})$$

$$+\,\lambda f(\hat{Y}+\frac{\delta-\varphi}{1-t_c})\,(1-t_c)$$

$$\cdot\,(\hat{Y}+\frac{\delta-\varphi}{1-t_c}-\bar{X}) - [\delta+\frac{d(\delta-\varphi)}{1-t_c}\,]$$

$$\cdot\,[\,(\frac{\hat{Y}+\frac{\delta-\varphi}{1-t_c}-\bar{X}}{\sigma})^2-1\,]\Big\} \gtrless 0$$

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Expectations and the Core Rate of Inflation

by Richard H. Jefferis, Jr.

Richard H. Jefferis, Jr., is a professor of finance in the School of Business Administration at the University of Colorado, Boulder, and was formerly a visiting scholar at the Federal Reserve Bank of Cleveland. The author would like to thank Michael Bryan, William Gavin, Jeffrey Hallman, and James Hamilton for helpful discussions and advice.

Introduction

Policymakers seeking to control inflation are confronted by a bewildering array of price statistics that often provide conflicting signals about the current inflation rate. The disparity among different measures of inflation is illustrated by figure 1, which depicts quarterly inflation rates implied by movements in several well-known price series between 1954 and 1987, and by table 1, which displays the correlation among inflation rates associated with a broader group of indices over the same period.[1] Although the CPI, the PCE deflator, and the PPI trend together, there is a wide variation in the movements of these price indices over periods as long as a quarter.

The discrepancy among inflation rates associated with different price indices has important implications for the conduct of monetary policy linked to inflation targets. If long-term increases in the price level are masked by statistical noise that is a consequence of changing circumstances in individual markets, then monetary policy linked to *any* index of current inflation will be affected by transient shocks as well as by the secular trend in prices. Although shocks to the price of individual commodities or groups of commodities do affect the cost of living, they do not necessarily reflect the impact of money growth on the price level. Nor is the appropriate policy response to these two types of inflation necessarily the same. Overall, both the source of noise and the amount of noise in individual price indices make them a poor choice for inflation targets.

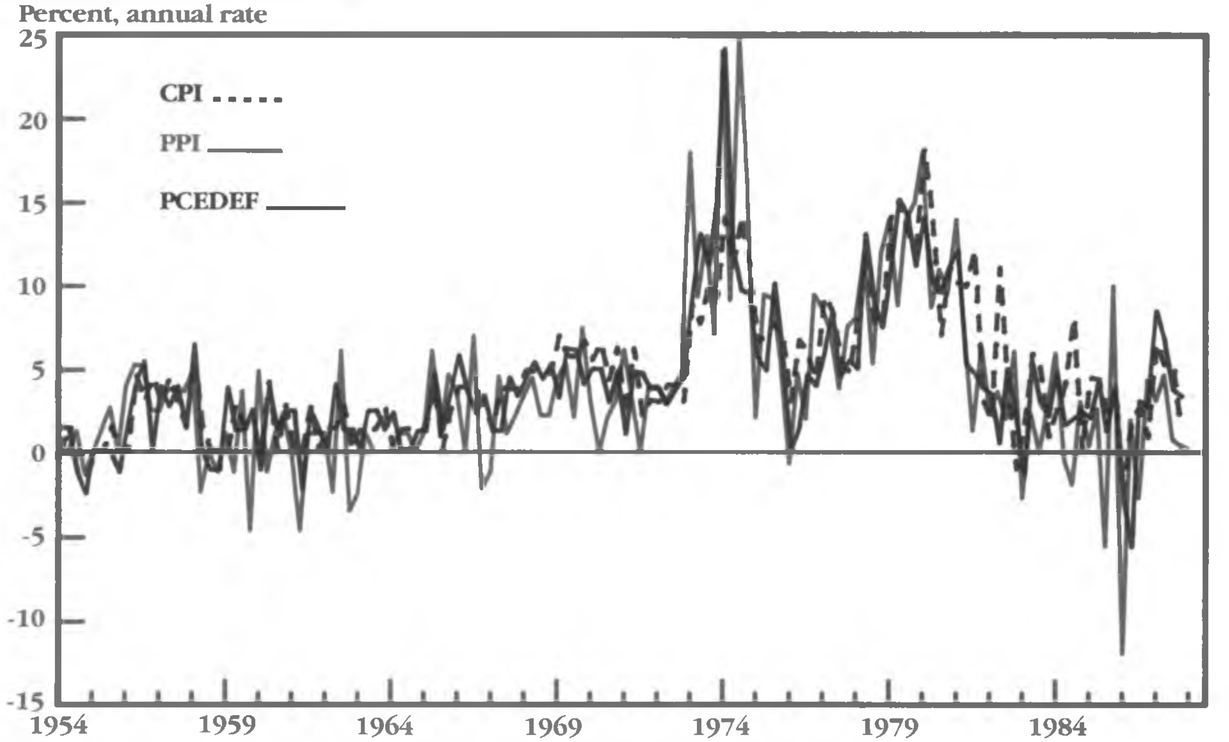
A related problem associated with using realized inflation as a guide for monetary policy is the timing of inflation signals. The inflation rate of last quarter or even last month is a poor guide for policy that seeks to influence the future course of the economy, yet this is the type of information provided by direct examination of any historical record. Forward-looking policies linked to a measure of current inflation should be based on what the past can tell us about the present and the future. To make the historical record useful, we need to extract from it information about current inflation and expected future inflation.

Inflationary expectations address both of these problems. Expectations are, by their nature, linked to long-term trends in the price level rather than to transient movements. They are forward-looking. Moreover, there is a remarkable degree of correlation among the inflation forecasts

■ 1 The price indices are the Consumer Price Index (CPI), the service component of the CPI (CPIS), the Producer Price Index (PPI), the PPI without food and energy (PPIWF), and the Personal Consumption Expenditure (PCE) deflator. Only the CPI, PPI, and PCE deflator appear in the figure.

FIGURE 1

Different Measures of Inflation, 1954-1987



SOURCES: U.S. Department of Commerce, Bureau of Economic Analysis, and U.S. Department of Labor, Bureau of Labor Statistics.

generated by different price indices and forecast methodologies. Realistic models of inflation discount current innovations in the inflation rate, which are largely noise, and focus instead on movements that tend to persist over time. As a result, different models of expected inflation agree on what is likely to occur in the immediate future, even when individual price series give conflicting signals about the current inflation rate. The common trend in the different series is an indicator of the pervasive price growth, or core inflation, that is of interest to policymakers.

I. Expected Inflation

We usually think of expected inflation as the expected rate of change in a particular price index, and judge different models of expectations by their ability to project movements in that index. The two criteria most commonly used to judge model performance are the mean squared error and bias of the inflation forecast. Both statistics are informative, since an unbiased forecast that fails to identify large, predictable movements in inflation will have a larger mean squared error than a forecast that is, on average, less accurate but better able to predict significant changes in the inflation rate.

Two types of statistical models used to forecast inflation have, in the past, performed equally well in terms of both bias and mean squared error.[2] Time series models identify temporal patterns in the inflation rate and use those patterns, combined with information about the recent history of inflation, to predict future inflation. These models are capable of identifying very complex relationships among inflation rates at different points in time, but tend to ignore other contemporaneous information that might be useful in forecasting.

Econometric models that incorporate information about interest rates or money growth attempt to remedy this shortcoming. Although the history of money growth is correlated with inflation, interest rates often take the place of money in forecasting models. The motivation for this choice is the notion that, in an efficient capital market, the

2 Fama and Gibbons (1984) compare pure time-series models and interest-rate models, and find that the interest-rate models yield a smaller root mean squared error in out-of-sample forecasts. The differences in forecast performance increase with the forecast horizon.

TABLE 1

The Correlation among Quarterly Inflation Rates Based on Different Price Indices between 1954 and 1987

	CPI	CPIS	PPI	PPIWF	PCEDEF
CPI	1.00	0.83	0.74	0.72	0.82
CPIS	0.83	1.00	0.48	0.56	0.56
PPI	0.74	0.48	1.00	0.79	0.76
PPIWF	0.72	0.56	0.79	1.00	0.78
PCEDEF	0.82	0.56	0.76	0.78	1.00

SOURCE: Author's calculations.

inflation premium in nominal interest rates is a sufficient statistic for expected inflation.

In practice, money may have some incremental predictive power, because the decomposition of nominal rates into an expected real return and an inflation premium is not observable, but is imposed on the data by the econometrician. To the extent that this decomposition is imperfect, the econometric model will fail to uncover the market's inflation forecast, even if movements in the nominal interest rate are completely determined by changes in the expected real rate and expected inflation, as theory would suggest. The merits of econometric models that extract inflation forecasts from interest rates and the empirical relevance of monetary growth for predicting inflation are issues that may be resolved only by examining the data.

II. Time Series Models

Time series models express current inflation as a weighted sum of past inflation and past changes in the inflation rate. The manner in which this history is translated into forecasts depends on the properties of the inflation process. When movements in the inflation rate tend to be transient, current innovations play a marginal role in the formation of expectations, and the historical record receives more emphasis in the inflation forecast. If, on the other hand, increases in inflation tend to persist, the inflation forecast will be closely linked to the behavior of prices during the recent past.

A time series model of inflation that has been found to forecast well is

(1) $I(t) - I(t-1) = \varepsilon(t) - \theta\varepsilon(t-1).$

In this model, $I(t)$ is the inflation rate at time t and $\varepsilon(t)$ is an impulse that affects that rate.[3] Conceptually, the impulse comes either from expansion of the money stock or from some change in market conditions, such as a drought or the threat of war in the Middle East. The current change in the inflation rate is determined by current and past impulses, where the weight assigned to the past is θ.

These models have an appealing interpretation in terms of expected and unexpected inflation.[4] From equation (1), we know that

(2) $\Delta I(t) = \varepsilon(t) - \theta\varepsilon(t-1).$

This implies that

(3) $\hat{I}(t) = I(t-1) - \theta\varepsilon(t-1)$

or that

(4) $\Delta I(t) = \Delta I(t-1) - \theta\Delta\varepsilon(t-1).$

Using the definition of $\Delta I(t)$ from equation (2) and the fact that $\Delta[\theta\varepsilon(t-1)] = \theta[\varepsilon(t-1) - \varepsilon(t-2)]$, we obtain

(5) $\Delta\hat{I}(t) = (1-\theta)\varepsilon(t-1).$

Expression (5) states that expected inflation follows a random walk, with an innovation variance that is $(1-\theta)^2$ times the variance of $\varepsilon(t)$. Values of θ close to 1 imply that most of the variance in inflation is accounted for by transient shocks, so that current innovations are not reflected in expected future inflation, while values of θ close to 0 imply that most of the variance is accounted for by movements in inflation that are expected to persist.[5]

Estimation of equation (1) for the different series described in table 1 and figure 1 yields values of θ that range from 0.45 for the PCE deflator to 0.70 for the PPI.[6] The evidence from the econometric model is therefore in accord with the intuition suggested by the data: A modest fraction of the quarterly innovation in inflation is reflected

3 The model described here is examined by Fama and Gibbons (1982).

4 Jeffrey Hallman suggested this interpretation.

5 Ansley (1980) provides an alternative interpretation that yields the same inference.

6 Maximum likelihood estimates are based on a sample of inflation rates from the first quarter of 1954 to the fourth quarter of 1987. The Breusch-Pagan Lagrange multiplier test for autoregressive conditional heteroscedasticity (ARCH) effects reveals that the data are conditionally heteroscedastic. All estimates involve an ARCH (2,0) model of the conditional variance, although this is found to have only a minimal impact on estimated parameter values and forecasts.

TABLE 2

The Correlation among Expected Quarterly Inflation Rates Generated by a One-Parameter Time Series Model. Inflation Is Assumed to Follow an IMA (1,1) Process

	CPI	CPIS	PPI	PPIWF	PCEDEF
CPI	1.00	0.92	0.93	0.90	0.90
CPIS	0.92	1.00	0.82	0.81	0.71
PPI	0.93	0.82	1.00	0.93	0.89
PPIWF	0.90	0.81	0.93	1.00	0.87
PCEDEF	0.90	0.71	0.89	0.87	1.00

SOURCE: Author's calculations.

in expected future inflation. In the case of the PCE deflator, a 1 percent increase in quarterly inflation is associated with a 0.55 percent increase in expected inflation. That fraction is 0.30 in the case of the PPI. An alternative perspective on the estimated value of θ is provided by examining the fraction of the variance of changes in quarterly inflation accounted for by changes in expected inflation. This number ranges from 10 percent in the case of the PPI to 30 percent in the case of the PCE deflator.

The effect of filtering the inflation-rate series with this model, and focusing on the expected inflation series implied by equation (3), is illustrated in table 2. The correlation among expected inflation rates inferred from the different price series is substantially greater than the correlation in realized inflation rates, even when expectations are generated by the parsimonious one-parameter time series model. For example, the correlation between the expected inflation rate inferred from the CPI and the expected inflation rate inferred from the PPI over 35 years of quarterly data is 0.93, while the correlation between the realized rates of inflation implied by these same indices is 0.74. Thus, the different price series yield highly correlated inflation forecasts, even though there is substantial disagreement about the current inflation rate among these series.

III. Econometric Models

Inflation forecasts based exclusively on the temporal pattern of past inflation ignore a great deal of potentially useful data. Information about money growth or interest rates will be without value only in the event that the history of inflation is a sufficient statistic for its expected future course. Both the tremendous amount of noise in the various inflation series and common sense suggest that this is unlikely.

Nominal interest rates are an especially appealing source of information, since the yield on fixed-rate debt instruments contains a premium that compensates the investor for expected depreciation in the purchasing power of money over the life of the instrument. The advantage of using interest rates to identify expected inflation, rather than modeling the link between money and prices directly, is that the inflation premium found in bond yields represents a consensus forecast of inflation over a fixed time interval known to the observer. In contrast, the history of money growth provides little information about when an increase in money will be reflected in prices, or even whether it will be reflected in prices rather than output. Focusing on bond yields rather than on money growth makes it unnecessary to consider the complex lag structures typical of macroeconomic models that attempt to characterize directly the link between money and prices.

Extracting inflationary expectations from bond yields is not a trivial exercise: Variations in nominal yields reflect changes in expected real returns as well as changes in expected inflation. (Yields may also contain a risk premium when inflation is uncertain, but this feature of returns is rarely modeled.) Neither component of nominal yields is observed directly, and models that exploit interest-rate data rely on auxiliary assumptions to separate expected real rates from expected inflation. The models discussed below are distinguished by the assumptions about the real rate process that are used to identify these components of the nominal interest rate.

One method of identifying the model is to assume that the expected real rate of return follows a random walk. This implies that

$$(6) \quad \hat{R}(t) = \hat{R}(t-1) + \zeta(t).$$

Then, if the realized real return is equal to the expected real return plus a noise term $\eta(t)$, the first difference of the observed real return takes the form

$$(7) \quad \Delta R(t) = \zeta(t) + \eta(t) - \eta(t-1).$$

TABLE 3

The Correlation among Expected Quarterly Inflation Rates when the Expected Real Rate Follows a Random Walk and the Nominal Yield Is the Sum of the Expected Real Rate and Expected Inflation

	CPI	CPIS	PPI	PPIWF	PCEDEF
CPI	1.00	0.95	0.96	0.93	0.92
CPIS	0.95	1.00	0.88	0.87	0.78
PPI	0.96	0.88	1.00	0.96	0.92
PPIWF	0.93	0.87	0.96	1.00	0.89
PCEDEF	0.92	0.78	0.92	0.89	1.00

SOURCE: Author's calculations.

TABLE 4

The Correlation among Expected Quarterly Inflation Rates Generated by a Regression-Based Model. The First Difference in Inflation Is Projected onto the First Difference in the 90-Day Treasury Yield

	CPI	CPIS	PPI	PPIWF	PCEDEF
CPI	1.00	0.95	0.94	0.93	0.91
CPIS	0.95	1.00	0.85	0.88	0.78
PPI	0.94	0.85	1.00	0.95	0.90
PPIWF	0.93	0.88	0.95	1.00	0.88
PCEDEF	0.91	0.78	0.90	0.88	1.00

SOURCE: Author's calculations.

This has a first-order moving average representation identical to that of equation (1). Estimation of this model yields an expected real return series.[7] Quarterly inflation forecasts are then constructed by subtracting the expected real return series corresponding to a particular price index from the yield on 90-day Treasury bills. The correlation among the inflation forecasts created in this manner is described in table 3.

The more sophisticated model of expectations yields inflation forecasts that are both more accurate and more highly correlated with each other than those from the time series model, even when the dynamics of the expected real interest rate are extremely simple.[8] The increased correlation is especially noticeable in situations where the correlation between the time series forecasts is lowest; for example, in the service component of the CPI and PPI. The high correlation among the fitted values from the interest-rate-based models suggests that all of the forecasts are tracking some underlying trend. The natural interpretation of that trend is the core rate of inflation.

This interpretation is reinforced by estimates from a closely related model. If expected real rates are constant or nearly constant between adjacent quarters, the main source of variation in Treasury yields is the inflation premium. This suggests a regression-based model of the form

$$(8) \qquad \Delta \pi (t) = \beta_0 + \Delta i(t) \beta_1 + \varepsilon (t),$$

where $\Delta \pi (t)$ is the change in inflation from one quarter to the next and $\Delta i(t)$ is the change in Treasury yields from the beginning of quarter $t-1$ to the beginning of quarter t. Estimation of this model indicates a statistically significant relationship between the change in Treasury yields and the change in inflation.[9]

The correlation among fitted values obtained by estimating equation (8) is documented in table 4. The strong resemblance between these results and those presented in table 3 suggests that whether interest rates are included in the model is a more important consideration than the manner in which they are incorporated. As before, the expected inflation forecasts track each other quite closely.

Adding lagged values of either the growth rate of money or the change in the growth rate of money to the regression equation has almost no impact on the fitted values for expected inflation, even though the regression coefficients associated with these variables are statistically

7 Application of the Breusch-Pagan test to the residuals from maximum likelihood estimates reveals ARCH effects. The figures in table 3 are based on fitted values from a maximum likelihood model where the conditional variance is ARCH (2,0). It is also worthwhile noting that the magnitude of the moving-average parameter is considerably less than in the results reported by Fama and Gibbons for monthly data. In other words, monthly data contain even more noise.

8 Fama and Gibbons (1984) document the superiority of this model relative to the time series model, using monthly data.

9 The model is estimated by maximum likelihood with an MA(1) error structure and an ARCH correction for conditional heteroscedasticity. The regresssion coefficient β_1 is statistically significant at 1 percent for all of the inflation series when the parameter covariance matrix is estimated from the information matrix, with or without the Newey-West correction for heteroscedasticity.

TABLE 5

The Correlation among Actual and Predicted Series for the CPI

	Actual	IMA (1,1)	Real rate is a random walk	Regression w/int. rates	Same w/int. rates and money
Actual	1.00	0.76	0.81	0.80	0.82
IMA (1,1)	0.76	1.00	0.96	0.97	0.95
Real rate is a random walk	0.81	0.96	1.00	0.99	0.98
Regression with interest rates	0.80	0.97	0.99	1.00	0.98
Same with interest rates and money	0.82	0.95	0.98	0.98	1.00

SOURCE: Author's calculations.

significant in all of the models. Indeed, the correlation among fitted values cannot be distinguished from the results presented in table 4. This is consistent with results reported by Fama (1982), who finds that interest rates contain most of the information about expected inflation that may be extracted from the history of money and output.

IV. Correlation among Forecasts from Different Methodologies

The results discussed above concern the correlation among the predicted values of different inflation series obtained with a specific econometric methodology. Inspection of the predicted values for a given series and different methodologies suggests that three observations are in order. First, the inflation forecasts from the different models are highly correlated; they appear to be tracking a common element. Second, the forecasts track each other more closely than they track actual inflation, consistent with my interpretation of the inflation series as signal plus noise. Third, the forecasts that incorporate interest-rate data are both more accurate than the forecasts generated by the time series model and more highly correlated with each other than with the time series model. Although table 5 describes the correlation among forecasts only for the CPI, similar results obtain for the other price series.

V. Hamilton's Model

A potential shortcoming of the econometric methodologies that I have considered is the extremely simple dynamics that are imposed on expected real interest rates and expected inflation in order to identify these components of the nominal rate process. Hamilton (1985) has proposed and estimated a model that permits richer dynamics in both components, and formalizes the intuition that the observed rate is equal to a signal (expected inflation) plus noise. The model, which contains the random-walk formulation (6) as a special case, assumes that the following relations among inflation, expected inflation, and real interest rates are stable over time:

$$(9)\quad \hat{r}(t) = k_r + \Phi(L)\,\hat{r}(t) + \Psi(L)\,\hat{\pi}(t) + \xi(L)\,\pi(t) + \varepsilon_r(t),$$

$$(10)\quad \hat{\pi}(t) = k_\pi + \alpha(L)\,\hat{r}(t) + \beta(L)\,\hat{\pi}(t) + \gamma(L)\,\pi(t) + \varepsilon_\pi(t),$$

$$(11)\quad \pi(t) = \hat{\pi}(t) + e(t).$$

Expected real rates and expected inflation are described by linear projections of these variables on their own past values and on the past values of actual inflation. The difference between expected inflation and actual inflation is a noise term, as in the simpler models discussed above. These assumptions, along with the assumption that the nominal rate is equal to the real rate plus the expected inflation rate, are sufficient to identify expected real rates and expected inflation. Note that equations (9) and (10), like equations (7) and (8), are statistical models of the relationships among these variables; there is no presumption that the lag polynomials $\Phi(L)$, $\psi(L)$, $\xi(L)$, $\alpha(L)$, $\beta(L)$, and $\gamma(L)$ represent the decision rules that agents use to form expectations about real rates and inflation.

Hamilton's model enjoys a second advantage relative to the simple models in addition to encompassing a wider variety of time series behavior. In equations (9), (10), and (11), the distinction between errors in expectations and errors that result from the econometrician's inability to observe expected real rates or expected inflation is modeled explicitly. The error terms ε_r and ε_π represent innovations in the expected real rate and expected inflation rate that are not captured by the linear projections of equations (9) and (10). These innovations arise because

TABLE 6

The Correlation among Expected Quarterly Inflation Rates Generated by Hamilton's Kalman Filter Model of Expected Inflation and Interest Rates

	CPI	CPIS	PPI	PPIWF	PCEDEF
CPI	1.00	0.57	0.40	0.57	0.60
CPIS	0.57	1.00	0.51	0.69	0.69
PPI	0.40	0.51	1.00	0.65	0.62
PPIWF	0.57	0.69	0.65	1.00	0.85
PCEDEF	0.60	0.69	0.62	0.85	1.00

SOURCE: Author's calculations.

we are unable to observe expectations. The error term $e(t)$ represents the difference between what agents thought would occur and what did in fact occur. Estimation of these parameters allows us to evaluate explicitly the contribution of these different sources of noise to the difference between expected inflation and actual inflation, making it unnecessary to assign an economic interpretation to the moving-average parameter in a time series model.

The estimated series are consistent with those produced by the other econometric models, in that innovations in the inflation rates appear to contain a substantial noise component.[10] One indicator of this phenomenon is the set of coefficients that represents the projection of expected inflation onto past values of inflation and expected inflation. In general, the sum of the coefficients for the four lagged values of expected inflation tends to be near one, while the sum of the coefficients for the four lagged values of actual inflation tends to be near zero. At the first two lags, the effect is even stronger; estimated parameter values imply that inflationary expectations tend to persist, while inflationary shocks tend to be reversed. This pattern, which is consistent with the time series properties of the errors in the simpler econometric models, is characteristic of all of the series except for the PCE deflator.[11] It suggests that expectations of inflation tend to persist, even in the face of significant changes in the current inflation rate.

A second indicator of the noise in the series for realized inflation is the fraction of the variation in the inflation rate accounted for by the expectation error series $e(t)$. This ranges from 20 percent in the case of the PCE deflator to 60 percent in the case of the PPI.

The expected inflation series from Hamilton's model differ from the estimates produced by the simpler econometric models in one important respect: The substantial increase in the number of explanatory variables yields a significant improvement in fit. As a result, the predicted values bear a stronger resemblance to the actual values and a weaker resemblance to each other. This fact is evidenced by the correlation among predicted values described in table 6.

VI. A Multiple Indicator Model

A multiple indicator model based on Hamilton's methodology incorporates the flexible dynamics of that model, but focuses on the common component of the different series rather than on the expected component of a particular series. Interest rates and a set of realized inflation series are driven by a single expected inflation series. This series is distinguished from the expected inflation series generated by Hamilton's model in that it provides information about pervasive price growth rather than about the behavior of a particular index.

I estimate the model by projecting expected inflation and the expected real interest rate onto their own past values and onto past values of the PPI. The realized values of the PCE deflator and the CPI both serve as indicators of the core rate. The realized value of inflation for each index is presumed to be equal to expected inflation plus a noise term.

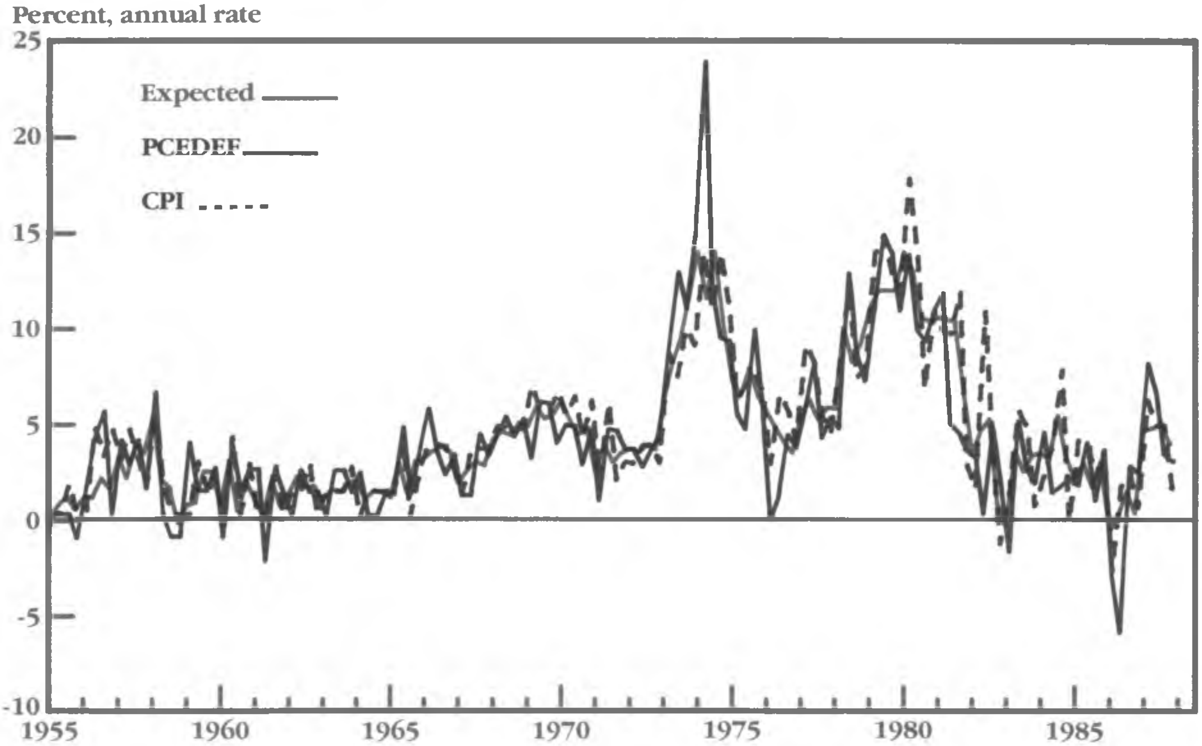
The expected inflation series for this model is presented in figure 2, along with the actual series for the CPI and the PCE deflator. Expected inflation exhibits the same time-series properties as do the individual series described above. Innovations in realized inflation are reflected only weakly in current expected inflation, which nonetheless displays a great deal of persistence.

10 I estimate the state space version of the model described in Burmeister, Wall, and Hamilton (1986). By doing so, I avoid dealing with the moving-average error terms that characterize the earlier formulation.

11 My estimates for the deflator series are qualitatively similar to those reported by Hamilton (1985) and Burmeister, Wall, and Hamilton (1986).

FIGURE 2

Expected and Realized Inflation, 1955 - 1987



SOURCES: U.S. Department of Commerce, Bureau of Economic Analysis, and U.S. Department of Labor, Bureau of Labor Statistics.

VII. Conclusion

Inflation targets may contribute significantly to the credibility of a monetary policy that is oriented toward controlling inflation. A potential problem with inflation targets is that inflexible rules would couple money growth to random shocks in the price level; the substantial noise in individual inflation series suggests that this concern is more than academic. Building flexibility into policy rules is one means of dealing with this problem, but flexibility tends to undermine the credibility of the commitment to control inflation. An inflation target that filters out these transient shocks, combined with a tight feedback rule from the filtered inflation rate to money growth, is an alternative that maintains credibility while mitigating the problems associated with noise in the policy targets.

Expected inflation is an indicator of the pervasive price growth, or core inflation, that interests the architects of monetary policy. The correlation among expected inflation rates from different price series and forecast methodologies suggests that these series are tracking the core rate. Signal extraction models formalize this intuition. Policy rules linked to the expected inflation series from any of the econometric models examined here are both forward-looking and reasonably insulated from index-specific shocks. Moreover, such broadly based targets would be difficult to manipulate. All of these properties suggest that expected inflation may serve as an effective guide to monetary policy.

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The Case of the Missing Interest Deductions: Will Tax Reform Increase U.S. Saving Rates?

by David Altig

David Altig is an assistant professor of business economics and public policy at Indiana University and is an economist at the Federal Reserve Bank of Cleveland. The author gratefully acknowledges helpful comments from Michael Bryan, Chris Carroll, Randall Eberts, and Jagadeesh Gokhale and excellent research assistance from Sharon Parrott.

Introduction

Beginning in tax year 1991, U.S. taxpayers may no longer deduct personal interest expense when calculating taxable income, thus completing the transition from the unlimited deductibility provisions that existed prior to enactment of the Tax Reform Act of 1986 (TRA86). In tax-speak, personal interest expense comprises interest payments not associated with mortgages on qualified residences or certain income-generating activities. Generally speaking, personal interest expense amounts to interest payments on consumer loans not secured by real estate.

Although a large share of household interest payments are associated with mortgage-related interest payments, which remain deductible under TRA86, disallowing deductions for personal interest expense is likely to have a substantial impact on consumer behavior.[1] Indeed, eliminating the deductibility of personal interest expense may, in the final analysis, be one of the more important legacies of TRA86.

It is certainly obvious that personal interest deductions had been increasingly exploited in the years preceding passage of TRA86. After trending upward during the 1950s, the growth of nonhousing interest deductions stabilized through the mid-1970s, fluctuating between 0.8 and 1.1 percent of GNP. After 1976, however, this percentage increased steadily, from 0.85 percent in 1977 to 1.7 percent in 1986 (see figure 1).

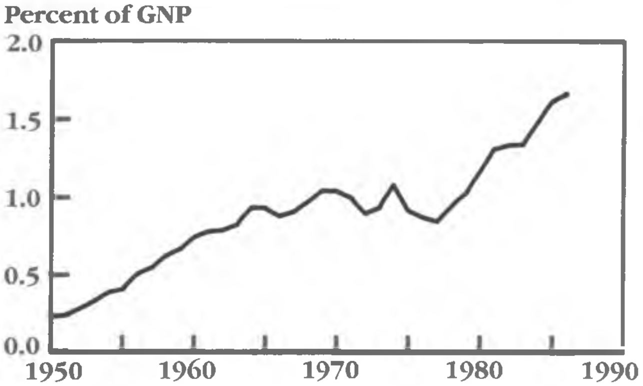
The period subsequent to 1976 was also distinguished by a downward trend in personal, private, and national saving rates (see figure 2). The coincidence of decreasing personal saving rates and increasing personal interest deductions can also be seen in figure 3, which plots personal saving (as a percentage of GNP) against nonhousing interest deductions (as a percentage of GNP).

While the negative relationship that appears in figure 3 does not necessarily imply that eliminating the deductibility of nonhousing interest

■ 1 The ratio of housing to nonhousing interest deductions on personal tax returns was 1.19 in 1966, 1.78 in 1976, and 1.78 again in 1986. The largest value of this ratio over the 1964-1986 period was 1.94, which was realized in 1983. Unfortunately, the Internal Revenue Service's *Statistics of Income*, from which these numbers are calculated, does not generally distinguish among the categories of nonhousing interest deductions. The nonhousing interest measures used in this paper therefore include interest expense associated with personal investment. Fortunately, available data suggest that investment interest expense claimed by individuals is small relative to personal interest expense. In 1977, for example, 65 percent of total household interest deductions were associated with home mortgages, 34 percent were associated with personal interest expense, and only 1 percent was associated with interest expense from investment activity.

FIGURE 1

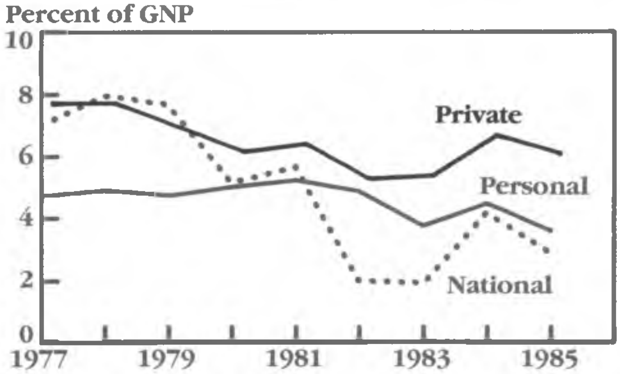
Nonhousing Interest Deductions



SOURCES: U.S. Department of Commerce, Bureau of Economic Analysis, and Internal Revenue Service.

FIGURE 2

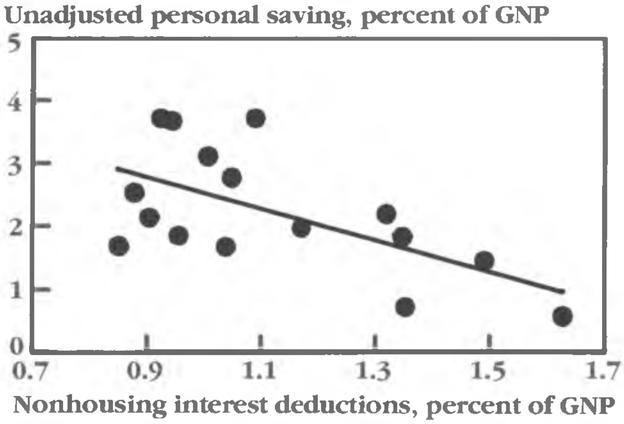
Saving Rates



SOURCE: Carroll and Summers (1987).

FIGURE 3

Personal Saving vs. Nonhousing Interest Deductions



SOURCES: U.S. Department of Commerce, Bureau of Economic Analysis, Internal Revenue Service, and Carroll and Summers (1987).

expense will cause an increase in the U.S. saving rate, it is commonly believed that removing incentives to dissave does indeed result in higher savings relative to income. To a large extent, this belief arises from the simple intuition that increasing the price of an activity—in this case, borrowing—will naturally lead to a decrease in that activity. Economic theory thus leads us to conclude that more restrictive tax treatment of personal interest expense will lead to less consumption and more saving.

Although empirical evidence is limited, it appears that the negative relationship between household borrowing subsidies and saving behavior suggested by economic theory can be found in real-world economies. Tanzi (1987) has shown that personal saving as a percentage of disposable income has tended to be lower in countries with the most generous tax treatment of personal interest expense (this evidence is also presented in Sheshinski [1990]). In a provocative comparison of U.S. and Canadian saving rates, Carroll and Summers (1987) argue that part of the historical divergence between observed saving rates in these two very similar economies is likely because, unlike taxpayers in the United States, Canadian taxpayers were unable to deduct personal interest expense.[2]

In this paper, I consider further some of the evidence presented by Tanzi and Carroll and Summers. Specifically, I ask two simple questions. First, do private saving rates tend to be higher, on average, in countries that prohibit the deductibility of personal interest expense? Second, do tax subsidies to borrowing help explain U.S.–Canadian saving rate differentials?

The empirical evidence I present gives affirmative answers to both questions. With respect to the first question, I examine private saving rates from 1975 to 1986 in a sample of 15 member countries of the Organisation for Economic Co-operation and Development (OECD). I find that private saving rates were indeed higher on average in countries without tax subsidies to consumption loans. These results confirm for private saving the observations made by Tanzi with respect to personal saving.[3]

■ 2 Limitations on interest deductions available to Canadian taxpayers also apply to interest expense from home mortgages. See the discussions in Carroll and Summers (1987) and Tanzi (1984).

■ 3 Private saving is the sum of saving by households, or personal saving, and saving by corporations.

Of course, simply comparing aggregated cross-country saving rates provides only casual evidence. Like the relationship in figure 3, such comparisons do not control for other causal factors. A more detailed analysis, which builds on the Carroll and Summers work, is provided in section III. The empirical models in this section add proxies for the U.S. subsidy rate on consumption loans to the Carroll and Summers regression equations for U.S.–Canadian saving differentials. The subsidy variables consistently appear with statistically and economically significant negative effects on private saving, a result that is remarkably robust across different specifications of the empirical model.

Even the more sophisticated analysis of section III has serious limitations — the data include only 24 annual observations, no attempt is made to control for simultaneity bias, and the subsidy proxies are admittedly crude, to name just a few. Furthermore, the effect of the borrowing subsidy variable is not consistently significant in regression models of the U.S. saving rate alone. Nonetheless, the results reported here are generally supportive of the assertion that consumption-loan subsidies may have important negative effects on saving behavior, and hence important implications for the long-run performance of the U.S. economy in the wake of TRA86.

I. A Simple Analytical Framework

Although the intuition for a negative relationship between favorable tax treatment of household borrowing and personal saving is readily apparent, introducing a simple analytical framework will help to organize the issues.

The framework presented here is a simple, perfect-certainty, overlapping generations model in which each generation lives three periods. Every generation consists of identical individuals who inelastically supply one unit of labor in the first two periods of life, retiring in the third. Utility is assumed to be a logarithmic, time-separable function of consumption given by

$$(1)\qquad U_t = \sum_{i=1}^{3} \beta^{i-1} \ln(C_{it}).$$

The variable β is the individual subjective time-discount factor, and the subscript t indexes each generation by date of birth.

Savers in the model have access to two types of assets: physical capital, denoted by a_{ij} for an age i individual of generation j, and private debt, which takes the form of consumption loans between generations.[4]

To make the model interesting, it is necessary that some generation chooses to borrow. I therefore assume that each generation is endowed with an identical, exogenous life-cycle labor productivity profile given by (ε_1, ε_2, 0), where ε_2 is sufficiently larger than ε_1 to ensure that the young always choose to borrow. Let borrowing by a young household born at time t be given by s_{1t}. Abstracting from population growth, market clearing in the consumption loans market requires that $s_{1t} = b_{2\,t-1}$, where $b_{2\,t-1}$ is lending by the generation that is middle-aged in time t.[5]

With these definitions in hand, the budget constraints for each generation are defined as

$$(2)\qquad C_{1t} = \varepsilon_1 w_t + s_{1t},$$

$$(3)\qquad C_{2t} = \varepsilon_2 w_{t+1} - [1 + r^d_{t+1}(1 - \delta_{t+1})]\, s_{1t} - a_{2t} - b_{2t},$$

and

$$(4)\qquad C_{3t} = (1 + r_{t+2})\, a_{2t} + [1 + r^d_{t+2}(1 - \rho_{t+2})]\, b_{2t},$$

where r is the rate of return to physical capital, r^d is the return to private debt, δ is the subsidy rate on borrowing (or, alternatively, the marginal tax rate on nonwage income for age 2 individuals), and ρ is the tax rate on interest income earned from the purchase of private debt. Equations (2), (3), and (4) embody the assumption that the young choose to borrow, the condition that all generations will consume their full lifetime resources (so that only middle-aged individuals save), and the simplifying assumption that the marginal tax rate on income from physical capital is zero.

Assuming interior solutions for individual saving and dissaving decisions, utility maximization implies the first-order conditions

4 Because the analysis here abstracts entirely from transaction costs, nothing essential is lost by ignoring the role of intermediaries and assuming that loan contracts are directly traded between generations.

5 The model abstracts from bequest motives and uncertainty, so all generations choose to "die" with no assets. The old will therefore never choose to accumulate capital or lend in the consumption loans market.

TABLE 1

Crowding-Out Effects of Increasing the Subsidy to Consumption Loans

Percentage Reduction in Steady-State Capital

δ	Benchmark	Population growth = 0	Productivity profile = (0, 10, 0)
0.01	0.5	0.6	0.8
0.02	1.0	1.1	1.6
0.03	1.6	1.7	2.4
0.04	2.1	2.3	3.3
0.05	2.6	2.9	4.2
0.06	3.2	3.5	5.0
0.07	3.7	4.1	6.0
0.08	4.3	4.7	6.9
0.09	4.8	5.3	7.8
0.10	5.4	6.0	8.8
0.11	5.9	6.7	9.8
0.12	6.5	7.3	10.8
0.13	7.0	8.0	11.9
0.14	7.6	8.7	12.9
0.15	8.2	9.4	14.0

NOTE: Each entry gives the percentage reduction in the steady-state capital stock when the subsidy rate on borrowing, δ, is increased from zero. The benchmark case assumes $\beta = 0.778$, $\theta = 0.25$, zero population growth, $(\varepsilon_1, \varepsilon_2, \varepsilon_3) = (1.5, 8.5, 0)$, and $\rho = 0.11$. The other cases maintain the benchmark assumptions, with the exception of the indicated parameters.
SOURCE: Author's calculations.

(5) $$C_{2t} = \beta\,[1 + r^d_{t+1}(1 - \delta_{t+1})]\,C_{1t},$$

(6) $$C_{3t} = \beta\,[1 + r^d_{t+2}(1 - \rho_{t+1})]\,C_{2t},$$

and

(7) $$C_{3t} = \beta\,(1 + r_{t+2})\,C_{2t}.$$

Equations (6) and (7) imply that, in asset-market equilibrium, $r_t = r^d_t(1 - \rho_t)$.

The long-run effect of changes in the subsidy variable δ can be demonstrated by a few simple simulation exercises. Table 1 reports the reduction in the steady-state capital stock caused by increasing the subsidy rate δ for particular parameterizations of the model.[6] In the benchmark case, which is described in table 1, increasing δ from 0 to 10 percent causes the steady-state capital stock to fall by 5.4 percent.[7] By extension, interest rates rise and per capita income falls.[8]

Table 1 also shows how factors that increase the demand for consumption loans amplify the crowding-out effects of allowing personal interest expense to be deducted for tax purposes. Thus, an increase in either the rate of population growth or the steepness of the productivity profile between young and middle ages results in larger percentage decreases in steady-state capital for a given change in δ. (See Bryan and Byrne [1990] and the references therein for a general discussion of the effects of demographics on aggregate saving in a life-cycle context.)

Because substantial disagreement persists among economists concerning the appropriate model of aggregate saving behavior, it is important to note that the qualitative results of the model presented here are not dependent on life-cycle assumptions. Altig and Davis (1989) show that changes in the subsidy rate on consumption loans can also have significant long-run negative effects on aggregate savings in models where parents and children are altruistically linked, as in Barro (1974). In fact, under the plausible assumption that the tax rate on interest income exceeds the subsidy rate on borrowing, Barro-type models predict that changes in subsidy rates can have large long-run effects on the size of the capital stock even when changes in the tax rate on interest income do not (see Altig and Davis [1989] for a full treatment of this issue).

II. Do Private Saving Rates Tend to Be Higher in Countries Without Borrowing Subsidies?

Table 2 answers this question directly. The answer is yes, at least for the subset of OECD countries examined here.[9] The results in table 2

6 The simulations reported in table 1 assume that all revenues raised (or lost) through distortionary taxation are rebated (or recovered) via lump-sum subsidies (or tax levies) to the affected generations.

7 In general, the direction of change in aggregate savings depends on the nature of the assumed preference structure. Under "standard" preferences, however, changes in the subsidy rate will have effects that are qualitatively the same as the ones reported here. The seminal discussion of this issue in an overlapping generations framework can be found in Diamond (1965).

8 The simulations reported in table 1 assume a Cobb-Douglas production technology, expressed in effective labor units as $y = k^{\theta}$. The steady-state rate of return to capital is therefore given by $\theta k^{\theta - 1}$. Thus, y is increasing in k, and r and r^d (by the asset-market clearing condition) are decreasing in k.

9 The countries are Australia, Austria, Belgium, Canada, Denmark, France, Ireland, Japan, the Netherlands, Norway, Sweden, Switzerland, the United Kingdom, the United States, and West Germany. The data are from *OECD National Accounts*.

TABLE 2

Average Private Saving Rates, 1975–1985

Group Averages

	Consumer interest not deductible	Consumer interest deductible
Simple average	10.68	8.65
Weighted average	11.14	8.38

Individual Country Averages

Consumer interest not deductible	Average saving rate
Australia	5.65
Austria	9.83
Belgium	12.73
Canada	11.81
France	9.74
Ireland	14.80
Japan	15.73
United Kingdom	7.33
West Germany	8.53
Consumer interest deductible	**Average saving rate**
Denmark	7.45
Netherlands	12.24
Norway	5.04
Sweden	5.23
Switzerland	13.72
United States	8.22

NOTE: Entries represent averages for subsets of 15 OECD countries. Countries are classified into deductible and nondeductible groups according to the information provided by Tanzi (1984). Weighted averages are constructed using within-group relative shares of real GDP. Real GDP figures are obtained from Summers and Heston (1988). Saving rates are expressed as percentages of GNP.
SOURCE: Organisation for Economic Co-operation and Development. *National Accounts of OECD Countries*, 1975-1987, Volume II.

were obtained by first averaging private saving as a percentage of gross domestic product (GDP) over the sample period 1975 to 1985 for each of the 15 countries considered.[10] The countries were then grouped according to whether tax subsidies were provided to interest expense from general (nonhousing) consumer credit.[11] Two sets of group-average measures are reported in table 2—one based on simple averaging and one obtained by weighting the individual country averages by within-group relative shares of real GDP.[12]

The average private saving rate for the sample period was 10.68 percent in countries without favorable tax treatment of personal interest expense and 8.65 percent in countries with favorable tax treatment of personal interest expense (11.14 percent and 8.38 percent, respectively, when country-specific saving rates are weighted by GDP shares). To put the magnitude of this difference in some perspective, the U.S. current account deficit was 5 percent of GDP in 1988. A 2 percent increase in the private saving rate for 1988 could therefore have financed more than one-third of the U.S. current account deficit, an amount equivalent to about $44 billion in 1988 dollars.

Table 2 also clearly shows that, in the chosen sample, average saving rates varied substantially among countries with similar tax treatment of personal interest expense.[13] It is impossible to know how much of the variation can be accounted for by economic, demographic, and policy variables without a more detailed investigation of the data. Unfortunately, the information that is necessary to conduct a more detailed

■ **10** The savings measures used here are net of depreciation. See Aghevli et al. (1990) for a general discussion of the OECD saving measures.

■ **11** Countries are classified into subsidy and nonsubsidy groups according to information reported in appendix III of Tanzi (1984). Updated information in Tanzi (1987) indicates that these classifications were still valid in 1985.

■ **12** Relative GDP shares are obtained using real GDP at international prices calculated by Summers and Heston (1988).

■ **13** There were also significant differences in the trend of saving rates for countries within the two groups. In the subsidy group, for instance, Norway, Sweden, and the United States experienced declining saving rates over the sample period, while Denmark, the Netherlands, and Switzerland all experienced fairly strong upward trends.

inquiry is difficult to come by.[14] Because of this difficulty, the balance of this paper focuses on a comparison between two countries for which data are more readily available: the United States and Canada.

III. Has the Subsidy Rate on Consumer Loans Reduced U.S. Saving Relative to Canadian Saving?

Following Carroll and Summers (1987), the starting point of the analysis in this section is a simple saving equation given by

$$(8) \quad S_t = \alpha_0 + \alpha_1\pi_1 + \alpha_2 UN_t + \alpha_3 SURP_t + \alpha_4 SHELT_t + \alpha_5 NW_t + \alpha_6 R_t^{at} + \eta_t,$$

where S_t is the time t differential between the U.S. and Canadian private saving rate (as a percent of GNP), π_t is the differential in inflation rates for consumer prices, UN_t is the differential in unemployment rates (as a percent of the total labor force), $SURP_t$ is the differential in net government saving (as a percent of GNP), $SHELT_t$ is the differential in the level of saving in tax-sheltered assets (as a percent of personal disposable income), NW_t is the differential in household net worth (as a percent of GNP), and R_t^{at} is the differential in weighted averages of after-tax returns to sheltered and nonsheltered saving.

Before proceeding to a discussion of my empirical work, it will be useful to introduce the rationale for including the particular regressors shown in equation (8). The inflation variable is included to control for the tendency of national income-account saving measures to overstate actual saving when inflation increases. The idea is that standard measures of income are distorted by changes in nominal interest rates that arise solely from changes in the rate of inflation or, more precisely, from the expected rate of inflation. This issue is examined in detail by Jump (1980). The expected sign of α_1 is positive if the type of measurement problem Jump identifies is the primary channel through which inflation rates help to explain aggregate savings.

The unemployment variable is a proxy for differences in cyclical conditions across the two countries. Assuming that changes in unemployment primarily reflect deviations from the equilibrium rate of unemployment, an appeal to the reasoning underlying the permanent-income hypothesis implies that $\alpha_2 < 0$. In other words, we expect higher unemployment and more dissaving when income is temporarily low.

The coefficient α_3 measures the relationship between public saving and private saving. In the simplest scenario, we expect to find $\alpha_3 = -1$ if the conditions necessary for Ricardian equivalence are true and $\alpha_3 > -1$ if those conditions are not true.[15] However, unambiguous predictions for the value of α_3 are complicated by the fact that equation (8) does not control for independent effects associated with government expenditures (see Aschauer [1985]).

The significance of the sheltered saving variable is the key finding of Carroll and Summers. $SHELT_t$ specifically measures the U.S.–Canadian differential in total personal saving in tax-sheltered forms (as a percentage of disposable personal income). In the United States, sheltered saving is represented by contributions to individual retirement accounts (IRAs). The Canadian equivalent of IRAs are registered retirement savings plans.

Carroll and Summers estimate values of α_4 that range between 1 and 2, implying that increases in the amount of saving in tax-sheltered assets are associated with greater than one-to-one increases in total private saving. Although this impact seems large, it is qualitatively consistent with microdata evidence presented by Venti and Wise (1987), who estimate that 80 to 90 percent of IRA contributions represent net increases in personal saving.

The final two variables, NW_t and R_t^{at}, are expected to enter equation (8) with negative and positive coefficients, respectively. The net-worth variable is included to capture the possibility that private saving, as measured on a national income accounts basis, changes as households seek to maintain target wealth-to-income ratios. Thus, as net worth rises relative to GNP, private saving tends to fall.

■ 14 I did examine many cross-sectional regressions with variations of the empirical specification employed by Feldstein (1980). In particular, I attempted to find whether this type of cross-sectional empirical saving model tends to underpredict the average private saving rate for countries without borrowing subsidies and overpredict the saving rate for countries with borrowing subsidies. For some of the models, I found regression errors were uniformly positive for the no-subsidy countries and uniformly negative for the countries with subsidies. However, the results were so sensitive to sample size, choice of regressors, and sample period that it was impossible to make a convincing case one way or the other. The general nonrobustness of Feldstein-like empirical saving models is also reported by Slemrod (1990) and Bosworth (1990).

■ 15 The literature on Ricardian equivalence is massive. Good general discussions can be found in Bernheim (1987, 1989) and Barro (1989a, 1989b).

TABLE 3

Regression Results

Coefficient Values	Model 1	2	3	4
CONST	−.027	.004	.004	.016
	(3.3)[a]	(.45)	(.43)	(2.3)[b]
INFL	.231	.197	−.017	.156
	(.95)	(1.1)	(.04)	(.43)
UN	.290	.506	.430	.404
	(.62)	(1.4)	(1.1)	(1.3)
SURP	−.837	−.267	−.281	−.365
	(3.4)[a]	(1.2)	(1.2)	(1.8)[c]
SHELT		1.74	1.98	−.665
		(4.2)[a]	(3.2)[a]	(.64)
R^{at}			−.254	.228
			(.55)	(.54)
NW				.179
				(2.9)[a]
Adj. R^2	.559	.760	.751	.824
ρ	.585	.408	.407	.226

a. The null hypothesis that the corresponding coefficient is zero can be rejected at the 99 percent confidence level.
b. The null hypothesis that the corresponding coefficient is zero can be rejected at the 95 percent confidence level.
c. The null hypothesis that the corresponding coefficient is zero can be rejected at the 90 percent confidence level.
NOTE: The dependent variable is the U.S.–Canadian differential in private saving relative to disposable income. All other variables are as defined in equation (8). The variable ρ is the first-order autocorrelation coefficient of the residual series. The numbers in parentheses are the absolute value of the *t* statistics for the corresponding coefficient estimate.
SOURCE: Author's calculations.

The after-tax real interest rate is included to capture the effects of changes in the return to saving. The expectation that $\alpha_6 > 0$ assumes that preferences cause substitution effects to dominate income effects and that ex post real rates are reasonable proxies for ex ante real rates.

An important consideration in discussing the expected signs of the coefficients in equation (8) is that I have described the relationships that would arise in an explicitly structural saving function. Equation (8) is, of course, decidedly nonstructural. Thus, coefficient estimates derived from regression analysis on equation (8) cannot be viewed as decisive indicators of the structural relationships between U.S.–Canadian saving differentials and the explanatory variables.[16] The appropriate interpretation of the approach taken here is that of an investigation into whether partial correlations of saving differentials and included regressors are consistent with structural–theoretical predictions.

Table 3 presents the results of several regressions based on equation (8). The data are annual and, with a few exceptions, are from Carroll and Summers (1987).[17] Model 1 in table 3 includes inflation, unemployment, and government surplus differentials as regressors. The coefficients on the inflation and government surplus variables have the anticipated sign, but only the government surplus variable is statistically significant.[18] The coefficient on the unemployment differential has the "wrong" sign, but is not statistically different from zero.

Models 2–4 in table 3 all include the differential in sheltered saving as a regressor. Models 2 and 3 essentially replicate the crucial Carroll and Summers result — the coefficient on *SHELT* is positive, large, and statistically significant. The coefficient on *SHELT* does become statistically insignificant when the U.S.–Canadian net wealth differential is added to the basic regression model.

■ 16 The problems in interpreting coefficient estimates from equation (8) are twofold. First, the coefficients in equation (8) are almost certainly "mongrel parameters," that is, unspecified functions of the underlying structural parameters. Second, no attempt is made to control for biases that may arise if the regressors are correlated with the error term η_t, a situation that seems likely. With respect to this latter problem, I did some limited experimentation with instrumental variables (IV) estimation. Unfortunately, the standard errors of the IV estimates were so large that no interesting inferences were possible.

■ 17 Unemployment rates are taken from the OECD *Labor Force Statistics*. The *SHELT* variable was constructed from data graciously provided by Chris Carroll (for Canada) and from data reported in Carroll and Summers (for the United States).

■ 18 The tables indicate coefficients that are statistically nonzero at the 90 percent, 95 percent, and 99 percent confidence levels. In the Carroll and Summers paper, reported *t* statistics are corrected for serial correlation. Although such corrections were made for all of the models reported in this paper, I have chosen not to report corrected *t* statistics for two reasons. First, almost all of the models estimated yield Durbin-Watson statistics that fall within the "inconclusive" range. Furthermore, although many of the models estimated yield "large" values of the first-order correlation coefficient of the residual series (ρ), the null hypothesis ρ = 0 is rarely rejected at the 95 percent confidence level. Second, work by Mishkin (1990) indicates that the type of correction employed by Carroll and Summers has undesirable properties in small samples. In most cases, the basic message is independent of whether *t* statistics are corrected or uncorrected.

TABLE 4

Regression Results Including Subsidy Variable

Coefficient Values	Model 5	6	7	8
CONST	.100	.092	.093	.064
	(6.1)[a]	(4.8)[a]	(4.7)[a]	(1.8)[b]
INFL	–.018	.004	.060	.108
	(.14)	(.03)	(.21)	(.36)
UN	–.077	.010	.025	.081
	(.32)	(.04)	(.09)	(.29)
SURP	–.260	–.210	–.206	–.247
	(1.8)[b]	(1.4)	(1.3)	(1.5)
SHELT		.353	.273	–.372
		(.89)	(.50)	(.44)
R^{at}			.069	.185
			(.21)	(.54)
NW				.062
				(1.0)
SUB	–4.25	–3.76	–3.80	–3.19
	(8.0)[a]	(4.9)[a]	(4.7)[a]	(3.2)[a]
Adj. R^2	.893	.891	.885	.885
ρ	–.128	–.150	–.170	–.232

a. The null hypothesis that the corresponding coefficient is zero can be rejected at the 99 percent confidence level.
b. The null hypothesis that the corresponding coefficient is zero can be rejected at the 90 percent confidence level.
NOTE: *SUB* is measured as the ratio of nonhousing personal interest deductions to adjusted gross income reported on itemized returns. See table 3 for other definitions.
SOURCE: Author's calculations.

Note also that the sign on the net wealth coefficient is positive and statistically significant.[19]

Table 4 presents results of regressions that add to models 1–4 a variable measuring the average borrowing subsidy. The subsidy variable is constructed as the ratio of total nonhousing interest deductions on personal tax returns to the adjusted gross income of all taxpayers with itemized deductions. This series on average subsidy rates is constructed from various issues of the *Statistics of Income for Individuals* (published by the Internal Revenue Service).[20]

The results in table 4 are striking. In every case, the null hypothesis that the subsidy variable has zero effect on private saving is easily rejected at the 99 percent confidence level. As would be expected, the explanatory power of the saving models also increases when the subsidy variable is included—in some cases, substantially.

It is necessary to bear in mind, however, that the average subsidy variable included in these regressions is at best a crude proxy for the variable that is theoretically important — namely, the *marginal* subsidy rate on consumption loans. In fact, it is difficult to distinguish movements in the subsidy variable that result from changes in tax incentives for borrowing from movements that result from shifts in the demand for consumption loans that are not associated with tax distortions.

For example, suppose that an individual, facing no change in borrowing subsidies, simply decides to borrow an extra $10 at the margin. Suppose further that the rate of interest on this loan is 10 percent. Then the individual's saving falls by $10 while his or her interest expense rises by $1. This single episode would suggest that the coefficient on the subsidy variable constructed from reported interest expense is –10, even though the borrowing behavior had nothing to do with tax-related borrowing subsidies.[21]

The regressions reported in table 5 replicate the regressions reported in table 4, with the subsidy variable calculated as 20 percent of the average nominal annualized return on three-month Treasury bills. Because personal interest deductibility provisions did not change during the sample period, exogenous changes in borrowing subsidies arose through two channels—changes in structural marginal tax rates and changes associated with variation in the rate of inflation in the context of a tax code that allowed for the deductibility of *nominal* interest expense. The subsidy variable used for the regressions in table 5 is designed to capture the effects of the latter channel.[22]

19 Carroll and Summers do not find the same sensitivity of the *SHELT* coefficient in their empirical analysis. The differences between their results and mine apparently result from the data. As subsequent results make clear, I find that no stable inference can be made about the relationship between U.S.–Canadian private saving differentials and differences in the amount of sheltered saving in the two countries.

20 Values for personal interest deductions are interpolated for the odd years from 1961-1971 and for 1974 by assuming that total interest deductions and mortgage-related interest deductions increase from the previous tax years at the same rate as total itemized deductions.

21 I am grateful to Chris Carroll for suggesting this example, as well as the alternative subsidy variable discussed in the subsequent paragraphs.

22 Twenty percent is chosen as a rough approximation to the average marginal subsidy rate on borrowing in accordance with the numbers reported in table 1 of Altig and Davis (1989).

TABLE 5

Regression Results with Alternative Subsidy Variable

Coefficient Values	Model 9	10	11	12
CONST	.016	.017	.017	–.012
	(2.1)[a]	(2.2)[a]	(2.1)[a]	(.72)
INFL	.227	.221	.142	.214
	(1.7)	(1.6)	(.43)	(.70)
UN	.072	.138	.113	.167
	(.28)	(.49)	(.37)	(.59)
SURP	–.506	–.438	–.441	–.458
	(3.5)[b]	(2.5)[a]	(2.4)[c]	(2.7)[a]
SHELT		.326	.428	–.873
		(.69)	(.69)	(1.0)
R^{at}			–.094	.171
			(.26)	(.48)
NW				.111
				(2.0)[c]
SUB	–2.94	–2.60	–2.58	–2.00
	(6.9)[b]	(3.9)[b]	(3.8)[b]	(2.9)[a]
Adj. R^2	.867	.864	.856	.877
ρ	–.074	–.068	–.064	–.232

a. The null hypothesis that the corresponding coefficient is zero can be rejected at the 95 percent confidence level.
b. The null hypothesis that the corresponding coefficient is zero can be rejected at the 99 percent confidence level.
c. The null hypothesis that the corresponding coefficient is zero can be rejected at the 90 percent confidence level.
NOTE: *SUB* is measured as 20 percent of the average annualized return on three-month Treasury bills. See table 3 for other definitions.
SOURCE: Author's calculations.

The results in table 5 do not differ appreciably from those reported in table 4. Although the coefficients on the subsidy variables decrease in magnitude, they remain large in absolute value and are always statistically different from zero. Furthermore, as in the regressions reported in table 4, inclusion of the subsidy variable renders the *SHELT* variable insignificant in all cases.[23]

Table 6 presents the results of regressions based on other variations of the model given in equation (8) for each of the two subsidy variables used in tables 4 and 5. Models 13–16 report the results of estimated models in which demographic and income-growth variables are included as explanatory variables, extensions suggested by the theoretical model in section I. Models 13 and 14 include the U.S.–Canadian differential in the percentage of the population aged 15-65. Models 15 and 16 report results in which the real GNP growth-rate differential is included as a regressor.[24]

Models 17 and 18 of table 6 report results with personal saving taken as the dependent variable and corporate saving introduced separately as a regressor. Analogous to the observations made about the government surplus variable in equation (8), corporate saving, after controlling for total wealth, should have a one-for-one negative effect on personal saving if individuals "pierce the corporate veil."[25]

In every case, including numerous regressions not reported in the tables, the result is the same. With the arguable exception of the government surplus variable, the borrowing subsidy, however measured, is the only explanatory variable that consistently shows up with a statistically significant effect on the U.S.–Canadian saving differential. Furthermore, the effect is always negative, and strongly so.

One further set of tests is reported in table 7. Because borrowing subsidies are zero for Canada, all variation in the subsidy variable arises from the U.S. data. The regressions in table 7 are therefore based on U.S. data alone.[26] Although the models with the subsidy variable constructed from Treasury bill rates yield results that are consistent with regressions based on U.S.–Canadian saving differentials, it is apparent

23 The subsidy proxy included in the table 5 regressions is, of course, subject to some of the same potential endogeneity problems as the subsidy variable employed in the table 4 regressions. For example, suppose that individuals in the economy anticipate better times ahead (and that these expectations are not closely related to effects that are controlled for by the inclusion of unemployment or GNP growth differentials). Permanent-income theory then tells us that the response will be an average increase in the desire to borrow. The resulting shift in the aggregate saving curve will drive up both real and nominal interest rates (holding expected inflation fixed).

24 If faster GNP growth means steeper life-cycle productivity profiles, the results of the simulations in section I suggest that coefficients on the GNP growth differential should be negative. However, the growth-rate differential may also pick up changes in cyclical conditions not captured by the unemployment-rate differential. This latter interpretation seems more likely in light of the significant positive coefficient estimates reported in table 6.

25 The necessity of controlling for total wealth is emphasized in the empirical studies by Auerbach and Hassett (1989) and Poterba (1989). The results in these papers suggest to me that individuals do indeed internalize corporate saving when making personal consumption decisions. However, the evidence is, as usual, ambiguous.

26 I am grateful to Randall Eberts for suggesting these regressions.

TABLE 6

Regression Results with Alternative Models

Coefficient Values	Model 13	14	15	16	17[a]	18[a]
CONST	.065	−.056	.082	−.057	.071	.015
	(1.8)[b]	(2.0)[b]	(2.3)[c]	(.30)	(3.9)[d]	(2.5)[c]
INFL	.111	.293	.271	.289	.263	.380
	(.36)	(.88)	(.90)	(.87)	(.99)	(1.3)
UN	.078	−.822	.294	.285	.102	.193
	(.27)	(1.4)	(.98)	(.84)	(.45)	(.80)
SURP	−.242	−1.26	−.306	−.481	−.192	−.439
	(1.4)	(2.9)[c]	(1.9)[b]	(2.7)[c]	(1.4)	(2.6)[c]
CORP					−.193	−.075
					(.74)	(.28)
SHELT	−.364	−2.4	.144	−.617	.64	.442
	(.42)	(2.0)[b]	(.16)	(.64)	(1.3)	(.70)
R^{at}	.186	.401	.283	.222	.034	.131
	(.52)	(.98)	(.85)	(.60)	(.14)	(.47)
NW	.059	.104	−.005	.084	−.014	.002
	(.87)	(1.6)	(.07)	(1.2)	(.34)	(.05)
POPRAT	.031	−.011				
	(.13)	(.05)				
YGROW			.204	.097		
			(1.6)	(.65)		
SUB 1	−3.19		−3.24		−2.65	
	(3.1)[d]		(3.4)[d]		(3.4)[d]	
SUB 2		−1.88		−1.89		−1.82
		(2.6)[c]		(2.6)[c]		(2.8)[c]
Adj. R^2	.878	.871	.924	.873	.910	.896
ρ	−.221	−.383	−.276	−.193	−.122	−.009

a. The dependent variable is the personal saving rate differential.
b. The null hypothesis that the corresponding coefficient is zero can be rejected at the 90 percent confidence level.
c. The null hypothesis that the corresponding coefficient is zero can be rejected at the 95 percent confidence level.
d. The null hypothesis that the corresponding coefficient is zero can be rejected at the 99 percent confidence level.

NOTE: *POPRAT* is the differential in the percentage of the population aged 15-65; *YGROW* is the differential in real GNP growth rates; *CORP* is the differential in private minus personal saving rates; *SUB* 1 is the subsidy variable as defined in table 4; and *SUB* 2 is the subsidy variable as defined in table 5. See previous tables for other definitions.

SOURCES: Author's calculations and *OECD National Accounts*, various issues.

TABLE 7

Regression Results for U.S. Personal Saving

Coefficient Values	Model 19	20	21	22	23	24
CONST	.060	.079	.107	.085	.066	.079
	(2.8)[a]	(8.1)[c]	(5.9)[c]	(8.5)[c]	(3.5)[c]	(8.2)[c]
INFL	–.675	.499	.028	.103	–.489	.325
	(2.8)[a]	(2.4)[a]	(.36)	(2.1)[b]	(2.2)[a]	(1.3)
UN	–.515	.086	–.071	.007	–.472	–.024
	(1.8)[b]	(.52)	(.24)	(.04)	(1.9)[b]	(.13)
SURP	–.271	–.643	–.106	–.373	–.444	–.636
	(1.3)	(3.2)[c]	(.42)	(2.3)[a]	(2.3)[a]	(3.2)[c]
SHELT	–1.21	–1.19	–.194	–.466	–1.52	–1.31
	(1.5)	(2.3)[a]	(.21)	(1.2)	(2.1)[a]	(2.6)[a]
R^{at}	–.771	.463			–.492	.322
	(3.0)[c]	(2.0)[b]			(2.0)[b]	(1.2)
NW	.002	.009	.006	.008	.001	.007
	(.54)	(3.2)[c]	(1.4)	(2.7)[a]	(.25)	(2.1)[a]
YGROW					.175	.090
					(2.6)[a]	(1.2)
SUB 1	2.80		–1.18		1.68	
	(1.8)[b]		(1.2)		(1.2)	
SUB 2		–3.44		–1.91		–2.60
		(4.0)[c]		(4.6)[c]		(2.4)[a]
Adj. R^2	.655	.791	.490	.756	.745	.797
ρ	–.195	–.012	.310	–0.18	.040	.052

a. The null hypothesis that the corresponding coefficient is zero can be rejected at the 95 percent confidence level.
b. The null hypothesis that the corresponding coefficient is zero can be rejected at the 90 percent confidence level.
c. The null hypothesis that the corresponding coefficient is zero can be rejected at the 99 percent confidence level.
NOTE: All variables refer to the U.S. values of the variables defined in earlier tables. The dependent variable is U.S. personal saving as a percentage of disposable income. See table 6 for other definitions.
SOURCE: Author's calculations.

that the effects of borrowing subsidies are far less consistent when included as regressors in the U.S. private saving-rate models. Note also that the sheltered saving variables are in some cases negative, large, and statistically significant. Explaining these anomalies is an important topic for future investigations.

IV. Concluding Remarks

The United States is not alone in recent attempts to mitigate the attractiveness of consumption loans through less-favorable tax treatment of personal interest expense. Recent tax reforms in Denmark and Sweden, for instance, have included provisions that effectively restrict the value of personal interest-expense deductions. Informative discussions of these changes and others can be found in Tanzi (1987) and Pechman (1988).

The evidence presented in this paper, though cursory by design, does indeed point toward important effects on aggregate saving behavior as a result of changes in the tax treatment of personal interest expense. In addition, as noted in section I, quite disparate models of intertemporal consumption behavior predict that changes in the degree to which consumption loans are subsidized through the tax system can have substantial effects on aggregate saving. The combination of these observations suggests that no assessment of U.S., or world, tax reform is complete without careful scrutiny of the treatment of personal interest expense.

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