Temporary Income Taxes and Consumer Spending

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Both economic theory and casual empirical observation of the U.S. economy suggest that spending propensities from temporary tax changes are smaller than those from permanent ones, but neither provides much guidance about the magnitude of this difference. This paper offers new empirical estimates of this difference and finds it to be quite substantial. The analysis is based on an amendment of the standard distributed lag version of the permanent income hypothesis that distinguishes temporary taxes from other income on the grounds that the former are "more transitory." This amendment, which is broadly consistent with rational expectations, leads to a nonlinear consumption function. Though the standard error is unavoidably large, the point estimate suggests that a temporary tax change is treated as a 50-50 blend of a normal income tax change and a pure windfall. Over a 1-year planning horizon, a temporary tax change is estimated to have only a little more than half the impact of a permanent tax change of equal magnitude, and a rebate is estimated to have only about 38 percent of the impact.

In 1968, faced with a classic case of demand inflation, Congress enacted a temporary increase in personal income tax payments to

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curb aggregate demand. In 1975, near the trough of our worst postwar recession, Congress enacted a tax rebate and other temporary decreases in taxes and increases in transfer payments designed to stimulate aggregate demand. Questions have been raised about the effectiveness of both measures.

The questions have both theoretical and empirical roots. On theoretical grounds, the permanent income–life cycle hypothesis seems to argue that temporary income tax changes should have little effect on consumer spending in principle. On empirical grounds, data on consumer behavior seem to suggest that the impacts of the two temporary taxes on spending were also small in practice. In 1968, the savings rate fell from 7.5 percent in the quarter immediately preceding the tax surcharge (1968:2) to only 5.6 percent in the first quarter of the surtax, suggesting that consumers kept spending despite the tax. In 1975, the savings rate ballooned from 6.4 percent just prior to the rebate to a stunning 9.7 percent in the quarter of the rebate (1975:2), suggesting that little of the rebate was spent.

The purpose of this paper is to study these two temporary tax changes in some detail. Precisely what prediction does economic theory make about the relative effectiveness of temporary versus permanent tax changes? And what conclusions can be reached from U.S. time-series data? The fact that the Carter administration asked for (but did not get) a repeat performance of the rebate in 1977 suggests that there is more than academic interest in the answers to these questions.

Section I outlines the theoretical issues, beginning with an idealized life-cycle model and proceeding to introduce some important "real world" considerations. Since the discussion shows quite clearly that the issue is an empirical one, Section II reviews previous empirical work on the subject very briefly. Section III explains the underlying basis of the empirical model of this paper, relating it to recent literature on rational expectations and the permanent income hypothesis (PIH), and then Sections IV and V show how this basic conceptual framework was converted into an operational empirical model. The estimates are presented and analyzed in Section VI, and Section VII summarizes the main conclusions.

I. The Implications of Theory: Pure and Impure

1. *The Pure Permanent Income–Life Cycle Theory*

As Eisner (1969) pointed out some time ago, the PIH casts doubt on the effectiveness of income tax changes that are labeled as temporary because such measures have only minor effects on permanent income.

To develop a theoretical benchmark for the marginal propensity to consume (MPC) that the PIH suggests might apply to a temporary tax, consider a rarefied world in which consumers with exogenous earnings streams select consumption paths to maximize lifetime utility. If capital markets are perfect, only the discounted present values of the earnings streams matter, so suppose all households earn a constant income y per year. Assume further that households differ only in age, a; that the real rate of interest is zero; and that the subjective rate of time discounting is also zero.[1] The question is, If income taxes are raised by z per capita for the period from $t = 0$ to $t = t_1$, how much less will consumers spend over this interval?

In answering this question, there are three population groups to keep track of. People who are "alive" (in the economic sense) at $t = 0$ and who live past the expiration of the tax suffer an income loss of $t_1 z$ over the period. If T denotes the length of life, then each such person of age a consumes a fraction $t_1/(T - a)$ of this loss during the years in which the temporary tax is in effect. Thus the change in consumption per capita is $\Delta C_1 = zt_1^2/(T - a)$.

Old people who are alive at $t = 0$, but die before $t = t_1$, lose only $(T - a)z$ in income. However, since they have MPCs of unity during the surtax period as a whole, their change in consumption per capita is $\Delta C_2 = (T - a)z$.

Finally, we must worry about people who are born between $t = 0$ and $t = t_1$. If a, a negative number between 0 and $-t_1$, denotes the age of such a person, and he lives for $t_1 + a < t_1$ years during the period $0 \leqslant t \leqslant t_1$, his income loss is $(t_1 + a)z$. Since he spends a fraction $(t_1 + a)/T$ of this income during the period $0 \leqslant t \leqslant t_1$, the change in his consumption is $\Delta C_3 = [z(t_1 + a)^2]/T$.

To derive the aggregate change in per capita consumption, weight these groups by the age distribution, considering the ages $-t_1 \leqslant a \leqslant T$. In the simplest case of a uniform age distribution, $f(a) = 1/T$, the change is

$$\Delta C = \int_0^{T-t_1} \frac{\Delta C_1}{T} da + \int_{T-t_1}^{T} \frac{\Delta C_2}{T} da + \int_{-t_1}^{0} \frac{\Delta C_3}{T} da.$$

Working out the integrals and dividing by the total income that is taxed away during the period ($t_1 z$ per capita), we obtain

$$\text{MPC} = \frac{t_1}{T}(\log T - \log t_1) + \left[1 - \frac{T}{2t_1} + \frac{(T - t_1)^2}{2Tt_1}\right] + \frac{t_1^2}{3T^2},$$

where the three terms show the contributions of the three different population groups.

[1] This is essentially the model introduced by Modigliani and Brumberg (1954).

To take a concrete example, suppose the typical lifetime of a household head as a household head is $T = 50$ years. Then, according to this formula, the MPC for a 1-year tax ($t_1 = 1$) is .09, while that for a 2-year tax ($t_1 = 2$) is .15.[2] The MPC = .09 for a 1-year temporary tax is only a rough benchmark representing the pure PIH, and there are a number of reasons why the theory probably systematically understates the responses of consumers to temporary taxes (see below), so we should not take this number too seriously. Still, there are two lessons worth drawing from this simple exercise—lessons that have often been forgotten in the temporary-tax debate.

a) Income gains and losses from temporary taxes will eventually be spent just like any other increment or decrement to lifetime resources: if less is spent at first (because $t_1 < T$), then more will be spent later. Thus if we want to inquire about the "effectiveness" of temporary taxes, we must specify a time horizon. Over a long enough run, they must be just as "effective" as permanent ones.

b) The so-called zero effect view—that consumers ignore the surtax and consume as if it never happened—does not represent the PIH at all. Instead, that theory says that consumers should spend precisely what they would on receipt of a windfall gain (or loss) of t_1z. In the illustrative calculation, this turns out to be the "9 percent effect" view.

2. *Caveats and Imperfections*

There are several reasons why surtaxes may affect spending more strongly than indicated by pure theory. First, tax-induced income changes that are not consumed must be saved. If windfall gains are used to purchase durable goods, consumer spending may rise much more strongly than consumption; the converse may happen when there are windfall losses. The magnitude of the marginal propensity to spend windfalls on durable goods is, of course, an empirical question.[3]

Second, some households may be subject to liquidity constraints that are usually ignored by the PIH. If we stay within the certainty context, these constrained households will react strongly to even temporary income changes.[4] Thus the aggregate MPC for a temporary tax is a weighted average of the low MPCs of unconstrained households and the high MPCs of constrained ones. Again, the importance of this phenomenon is an empirical question.

[2] By way of comparison, as $t_1 \rightarrow T$, the MPC $\rightarrow$ 5/6.

[3] See, e.g., Darby (1972).

[4] See Blinder (1976) and Dolde (1978). For a look at one particular type of uncertainty, see Foley and Hellwig (1975), which shows that this result may not carry through to the uncertainty case.

Third, as Okun (1971) pointed out, consumer behavior depends on what people believe rather than on what the government announces. If consumers disbelieve the government when it tells them that a tax hike is only temporary, then the spending response will be greater than that suggested by a naive application of the PIH.[5] Since the perceived duration of the surtax, not the declared duration, is relevant from the standpoint of the PIH, this too raises an empirical issue.

Finally, we must recognize the possibility that households may not do the kind of rational long-term planning envisioned by Modigliani and Brumberg (1954) and Friedman (1957) or, what amounts to the same thing, have very high subjective discount rates. If they are very shortsighted, then temporary fluctuations in disposable income may have substantial effects on spending.

II. Previous Empirical Work

Okun's (1971) study opened the empirical debate on this issue. Using the consumption equations of four econometric models, he compared the "full effect" view that the 1968 surtax was just as effective as a permanent tax increase to the "zero effect" view that consumers totally ignored the surtax. While he concluded that the full effect view fit the data better, an intermediate "50 percent effect" view actually does better than either extreme.[6] Springer (1975) criticized Okun's econometric procedures and then performed a similar experiment with a consumption function based on the PIH. He concluded that the zero effect view performed better.

Juster (1977), using a series of savings equations based on the Houthakker-Taylor (1966) model, reached conclusions about the 1975 rebate similar to those of Okun for the 1968 surtax. But Modigliani and Steindel (1977) found that the rebate had very little impact over a horizon of 1 or 2 quarters. Their modified version of the life-cycle model implied, however, a virtually full effect view over a 6-quarter horizon. Modigliani and Steindel assumed that the nonrebate portions of the 1975 tax cuts were treated like permanent taxes and handled the 1968 surtax with dummy variables.

The existing empirical literature thus offers little consensus. The issue seems quite open.

III. The Distributed Lag Model of Consumption

The basic vehicle for investigating the effectiveness of temporary income taxes in this paper is the distributed lag version of the PIH.

[5] This point has rather less cogency with respect to tax cuts; but here too consumers may believe them to be more permanent than the government announces.

[6] On this, see Blinder and Solow (1974, pp. 107–9).

While this is the standard way of implementing the PIH empirically, the recent literature on rational expectations has seemed to raise doubts about its validity.[7] This section shows how the PIH and rational expectations together lead to an estimating equation very much like the one I use in this paper.

As Muth (1960) pointed out, the PIH is basically a *forward-looking* model of consumer behavior. It states that consumers, in deciding on their current spending, weigh their current asset holdings, their current income from labor, and their expected future income from labor. Specifically, permanent income at time t is defined as

$$Y_t^p = A_t + \sum_{s=0}^{\infty} \frac{{}_tY_{t+s}}{(1 + r)^s}, \tag{1}$$

where A_t is the stock of real assets at the beginning of period t, Y_t is noninterest income in period t, and ${}_tY_{t+s}$ is the mathematical (i.e., rational) expectation of Y_{t+s} that is formed at time t. (By convention, ${}_tY_t = Y_t$.) However, if the stochastic process generating income can be described by a time-series model such as

$$Y_t = a_1Y_{t-1} + a_2Y_{t-2} + \ldots + a_{n+1}Y_{t-n-1} + \epsilon_t, \tag{2}$$

where ϵ_t is a white noise error term, then the resulting empirical model of consumption will be *backward looking*. For example, if the theoretical consumption function is

$$C_t = \delta + kY_t^p + u_t, \tag{3}$$

then the empirical consumption function will be

$$C_t = \delta + kA_t + b_0Y_t + b_1Y_{t-1} + \ldots + b_nY_{t-n} + u_t. \tag{4}$$

To see this, it is only necessary to note that the (rational) expectation of income in period $t + s$ can, in view of (2), be based only on the information set $\{Y_t, Y_{t-1}, Y_{t-2}, \ldots\}$. Thus, for example,

$$\begin{aligned} {}_tY_{t+1} &= a_1Y_t + a_2Y_{t-1} + a_3Y_{t-2} + \ldots + a_{n+1}Y_{t-n}, \\ {}_tY_{t+2} &= a_1 {}_tY_{t+1} + a_2Y_t + a_3Y_{t-1} + \ldots + a_{n+1}Y_{t-n+1}, \\ &= (a_1^2 + a_2)Y_t + (a_1a_2 + a_3)Y_{t-1} + \ldots, \end{aligned}$$

and so on. Substituting all such expressions into the definition (1) and then into the consumption function (3), it is clear that (4) is derived. As pointed out by Sargent (1978) and others, the coefficients b_i in equation (4) will be complicated functions of the coefficients a_i in (2).

Suppose then that, as suggested by Dolde (1976), we can distinguish among two or more sources of income whose generating functions (2) may differ. The PIH in conjunction with rational expectations then

[7] See Lucas (1976) and esp. Hall (1978).

implies that the b's in (4) should follow a different pattern for each income source. A simple example will illustrate this point and also give us some feeling for possible magnitudes. Consider several income sources, each of which is generated by a first-order autoregressive:

$$Y_{it} = \rho_i Y_{i,t-1} + \epsilon_{it}. \tag{5}$$

Working out the expectations and plugging into (1) gives us a simple expression for the permanent income attributable to each source:

$$Y_{it}^p = \frac{1 + r}{1 + r - \rho_i} Y_{it}, \quad \text{if } \rho < 1 + r. \tag{6}$$

Notice that, despite the long time horizon contemplated by the PIH, consumption depends only on current income. In general, consumption will depend on past income only up to lag n, where $n + 1$ is the longest lag considered in equation (2).

Now compare two income sources, one of which is entirely permanent ($\rho = 1$) and the other of which is entirely transitory ($\rho = 0$). A \$1.00 increase in the permanent component will, according to (6), raise permanent income by $\$(1 + r)/r$ and thus raise consumption by $\$k(1 + r)/r$. This may imply a very large immediate spending response.[8] By contrast, a \$1.00 fluctuation in the purely transitory component will, again according to (6), raise permanent income by only \$1.00 and thus raise consumption by only $\$k$. The lesson, of course, generalizes and applies far beyond the confines of first-order autoregressives: *income sources deemed to be more permanent will elicit prompter spending responses than income sources deemed to be more temporary.* The application of this principle to permanent versus temporary changes in taxes is apparent and immediate and was elucidated clearly by Lucas (1976). It is the basic notion underlying the empirical model to be developed in the next section.

However, lest confusion arise, I should stress that there is no sense in which the rationality of expectations is either assumed or imposed in the consumption functions estimated here. My point is only that the distributed lag formulation of the PIH is consistent with rational expectations. As Sargent (1978) has emphasized, rational expectations delivers a set of restrictions across equations (2) and (4) that can be imposed in estimating the two jointly. I have made no attempt to impose these restrictions here because my interest was in getting the best possible consumption-function estimates, not in testing rationality. Furthermore, it is well known that quite different models

[8] Suppose the rate of subjective time discounting is equal to the rate of interest, so that a constant consumption stream is optimal, and that B is the lifetime propensity to consume (i.e., $1 - B$ is the propensity to bequeath). Then k will be $Br/(1 + r)$, so that $k(1 + r)/r$ will be B, which is close to unity.

of consumption behavior (e.g., habit persistence) can lead to an estimating equation very much like (4). It is not my purpose to discriminate among alternative ways of arriving at (4).

IV. Derivation of an Estimating Equation

The preceding discussion makes it clear that different distributed lag coefficients might be associated with different sources of income. While the actual empirical analysis considered four types of income, the model is most readily explained if I suppose there are only two: income (positive or negative) attributable to temporary tax measures, which I denote as S_t ("special income"); and all other disposable income ("regular income"), which I denote as R_t. The R_t should not be confused with permanent income, since it has both permanent and transitory components. The basic idea underlying the estimating equation is that S_t is identifiably "less permanent" than R_t.

Suppose consumption responds to R_t according to a set of distributed lag weights: $w_j = \partial C_t / \partial R_{t-j}, j = 0, 1, \ldots, n$. Since the w_j depend on the stochastic process generating R_t, it is worth reporting that the deviations of R_t from a logarithmic time trend are well described by the following second-order autoregressive:[9]

$$y_t = \underset{(.09)}{1.28} y_{t-1} - \underset{(.09)}{.35} y_{t-2}, \qquad R^2 = .91, \text{ D-W} = 1.86.$$

When income follows a second-order autoregressive, permanent income as defined in (1) is:

$$Y_t^p = A_t + K_t + \frac{(1+r)^2}{(1+r)(1+r-a_1) - a_2} y_t + \frac{a_2(1+r)}{(1+r)(1+r-a_1) - a_2} y_{t-1},$$

if $a_1 + a_2/(1+r) < 1 + r$, where K_t is the present value of the trend component of labor income and y_t and y_{t-1} are current and lagged deviations from trend. Given the estimates of a_1 and a_2 above, and for $r = .0074$ (a 3 percent annual real interest rate), the implied coefficients are $Y_t^p = A_t + K_t + 13.5y_t - 4.7y_{t-1}$. This leads us to expect a very large value of w_0, followed by swiftly declining w's—possibly even turning negative. The empirical results bear this out.

As Lucas (1976) has argued, income changes that are clearly "more temporary" than regular income should get different spending

[9] Standard errors are in parentheses. Longer autoregressives, however, give slightly better fits. E.g., if $t - 11$ is the longest lag allowed to enter the regression, significant coefficients are obtained at lags 1, 2, 3, 4, and 11. An F-test for the zero restrictions implied by the second-order model, however, yields an F-ratio of only 1.77, which is well below the critical 5 percent point of the χ_9^2 distribution.

coefficients. To develop a model of the distributed lag response of C_t to S_t, first break down S_t into its components:

$$S_t = S_t^1 + S_t^2 + \ldots + S_t^m, \tag{7}$$

where S_t^i indicates the income gain or loss in quarter t from the ith temporary tax. (In the empirical work, $m = 3$.) It will help clarify the treatment of the S_t^i if I define a hypothetical set of lag coefficients β_j as the effect on C_t of a \$1.00 pure windfall gain received in quarter $t - j$.

The treatment of S_t^i depends on whether or not the ith temporary tax is still in effect. If it is, I assume that S_t^i is treated as a weighted average of regular and windfall income, so that it gets the distributed lag weights

$$\gamma_j = \frac{\partial C_t}{\partial S_{t-j}^i} = \lambda w_j + (1 - \lambda)\beta_j, \qquad j = 0, 1, \ldots, n \quad 0 \leq \lambda \leq 1. \tag{8}$$

If the temporary tax is no longer on the books, I assume that consumers look upon S_{t-j}^i in retrospect as if it had been a pure windfall and so apply the distributed lag coefficients β_j. By introducing a dummy variable defined as:

$$D_t^i = 1 \text{ if the } i\text{th temporary tax remains in force in quarter } t,$$
$$= 0 \text{ otherwise,}$$

it is possible to combine these two hypotheses into a single expression:

$$\gamma_j^i(t) = D_t^i[\lambda w_j + (1 - \lambda)\beta_j] + (1 - D_t^i)\beta_j, \tag{9}$$

where the notation now indicates that the γ weights depend both on calendar time and on the specific tax under consideration (because of the dummy variable).

An interesting point arises here. Standard pre-rational-expectations approaches to consumption-function estimation would suggest that $\Sigma\beta_j$ and $\Sigma\gamma_j$ be constrained to equal Σw_j, apparently meaning that the "long-run MPC" out of any type of income is identical. However, the PIH-cum-rational-expectations approach suggests no such adding-up constraint. To see this, follow Sargent (1978, pp. 681–82) in rewriting (2) in the form $X_t = HX_{t-1} + \eta_t$ (Sargent's eq. 8), where

$$X_t = \begin{bmatrix} Y_t \\ Y_{t-1} \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ Y_{t-n} \end{bmatrix}, \quad H = \begin{bmatrix} a_1 & a_2 & \ldots & a_n & a_{n+1} \\ 1 & 0 & & 0 & 0 \\ 0 & 1 & & \cdot & \cdot \\ \cdot & 0 & & \cdot & \cdot \\ \cdot & \cdot & & \cdot & \cdot \\ \cdot & \cdot & & 0 & \cdot \\ 0 & 0 & & 1 & 0 \end{bmatrix}, \quad \eta_t = \begin{bmatrix} \epsilon_t \\ 0 \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ 0 \end{bmatrix},$$

and $Y_t = dX_t$, where $d = (1, 0, \ldots, 0)$. As Sargent notes (his eq. 9), rational expectations implies ${}_tX_{t+s} = H^sX_t$, whence ${}_tY_{t+s} = dH^sX_t$. Substituting this into (1) gives the following expression for permanent income:

$$Y_t^p = A_t + \left[\sum_{s=0}^{\infty} \frac{dH^s}{(1+r)^s}\right] X_t,$$

which is of the form (4) with coefficients

$$b \equiv (b_0, b_1, \ldots, b_n) = \sum_{s=0}^{\infty} \frac{dH^s}{(1+r)^s}.$$

The sum of the b_j has no obvious interpretation. Thus, in a model with several sources of income, there is no particular reason why the various sets of distributed lag coefficients should have a common sum.

Where, then, does the lifetime budget constraint enter? The answer is that (4) implies a unitary lifetime MPC for any values of k and the b_j. The proof involves some straightforward but tedious algebraic manipulations of the difference equations (4) and

$$A_{t+1} = (1+r)A_t + Y_t - C_t \tag{10}$$

and hence is relegated to Appendix A.

With these preliminaries out of the way, it is easy to explain the estimating equation. If there were no special taxes to worry about, the basic empirical model of consumer behavior would be as follows:

$$\begin{aligned} C_t = k_0 + k_1 r_t Y_t + \sum_{j=0}^{n} w_j Y_{t-j} \\ + k_2(W_t - A_t) + \sum_{j=0}^{q} k_{3+j} A_{t-j} + u_t, \end{aligned} \tag{11}$$

where r_t is the rate of interest, W_t is consumer net worth at the beginning of period t, and A_t is the market value of stock market wealth at the beginning of period t. The specific way in which assets are entered into the consumption function, including the constraint that $k_3 + k_4 + \ldots + k_{3+q} = k_2$, is suggested by the MIT-Penn-SSRC (MPS) model and is unimportant to what follows.

Now consider the separation of disposable income into its two components:

$$Y_t = R_t + S_t. \tag{12}$$

The way I have defined the γ's means that (11) is expanded to:

$$C_t = k_0 + k_1 r_t Y_t + k_2(W_t - A_t) + \sum_{j=0}^{q} k_{3+j} A_{t-j}$$

$$+ \sum_{j=0}^{n} w_j R_{t-j} + \sum_{j=0}^{n} \gamma_j^1(t) S_{t-j}^1 + \ldots + \sum_{j=0}^{n} \gamma_j^m(t) S_{t-j}^m + u_t. \tag{13}$$

Substituting (9) and (12) into (13), and rearranging terms, gives:

$$\begin{aligned} C_t = k_0 + k_1 r_t Y_t + k_2(W_t - A_t) + \sum_{j=0}^{q} k_{3+j} A_{t-j} \\ + \sum_{j=0}^{n} w_j (Y_{t-j} + \lambda X_t^j - S_{t-j}) \\ + \sum_{j=0}^{n} \beta_j \left[(1 - \lambda) X_t^j + \sum_{i=1}^{m} (1 - D_t^i) S_{t-j}^i \right] + u_t, \end{aligned} \tag{14}$$

where $X_t^j \equiv D_t^1 S_{t-j}^1 + \ldots + D_t^m S_{t-j}^m, j = 0, \ldots, n$. This is not the actual estimating equation because additional income sources were distinguished, because the distributed lag coefficients were constrained in several ways, and because corrections were made for both heteroscedasticity and serial correlation in the error term. Details are spelled out in Appendix B. Nonetheless (14) is the most useful form for interpreting the estimated parameters. The model is nonlinear because of the parameter λ—the crucial parameter of this study.

As noted above, theory does not imply that $\Sigma w_j = \Sigma \beta_j$. To investigate this further, this adding-up constraint was imposed in (14) and its validity tested as follows.[10] Let ℓ denote the likelihood ratio. Then, under the assumption of normality, $-2 \log \ell = T \log (\mathrm{SSR}_r/\mathrm{SSR})$ is distributed as a χ^2 with r degrees of freedom, where: T = number of observations (= 100 throughout this paper), SSR = minimized value of the sum of squared residuals in the unconstrained regression (eq. [14] in this case), SSR_r = minimized value of the sum of squared residuals in the constrained regression (obtained by imposing $\Sigma\beta = \Sigma w$), and r = number of restrictions (= 1 in this case). As reported in table 1, row 1, the constraint was rejected at the 10 percent level but not at the 5 percent level. There being no persuasive theoretical rationale for it, the constraint was dropped.

Tacitly, however, (14) embodies a number of other constraints that are equally lacking in theoretical justification—constraints that each type of regular income is subject to the same set of distributed lag coefficients. These constraints were tested by a series of likelihood ratio tests, which are described in the balance of this section.

[10] See Goldfeld and Quandt (1972, p. 74).

TABLE 1

χ^2 TESTS OF CONSTRAINTS

Unconstrained Model	Constraint Tested	df	Test Statistic
1. Eq. (14)	$\Sigma w = \Sigma\beta$	1	3.67
2. Eq. (16)	$v_j = w_j$ for all j	3	12.02
3. Eq. (16)	$\Sigma v_j = \Sigma w_j$	1	Approximately 0*
4. Eq. (18) with $\Sigma v = \Sigma w$	$\phi_j = w_j$ for all j	3	19.02
5. Eq. (18) with $\Sigma v = \Sigma w$	$\Sigma\phi_j = \Sigma w_j$	1	12.52
6. Eq. (18) with $\Sigma v = \Sigma w$	$\lambda = 0$	1	1.32
7. Eq. (18) with $\Sigma v = \Sigma w$	$\lambda = 1$	1	2.84
8. Eq. (18) with $\Sigma v = \Sigma w$	$k_2 = 0$	1	17.81

NOTE.—Critical levels for the χ^2 distribution are:

df	10% Point	5% Point	1% Point
1	2.71	3.84	6.63
3	6.25	7.81	11.34

*Due to rounding error, the actual computed test statistic was slightly negative. When this constraint was tested in the context of eq. (18), it produced a test statistic of 1.32.

First, "regular" disposable income was disaggregated into its two main components—personal income and "regular" personal taxes:

$$R_t = P_t - T_t.^{11} \tag{15}$$

Personal income is assumed to be spent according to the lag weights w_j, while regular taxes are assumed to be subject to a different set of lag weights v_j. Special taxes are treated as previously explained, except that the v's replace the w's in equations (8) and (9). That is, while the tax is on the books, a special tax is treated as a weighted average of a permanent tax and a windfall. Thus the basic consumption function becomes:

$$\begin{aligned} C_t = k_0 + k_1 r_t Y_t + k_2(W_t - A_t) + \sum_{j=0}^{q} k_{3+j} A_{t-j} \\ + \sum_{j=0}^{n} w_j P_{t-j} - \sum_{j=0}^{n} v_j (T_{t-j} - \lambda X_t^j) \\ + \sum_{j=0}^{n} \beta_j \left[(1-\lambda) X_t^j + \sum_{i=1}^{m} (1 - D_t^i) S_{t-j}^i \right] + u_t. \end{aligned} \tag{16}$$

Since (as explained below) the distributed lag coefficients are constrained to follow a third-degree polynomial with a zero end-point

[11] In making this separation, I departed a bit from national income accounting conventions by including the employer's share of social insurance contributions in both P and T.

constraint, the null hypothesis that the v_j are equal to the w_j imposes three constraints on equation (16). Row 2 of table 1 shows that these constraints were resoundingly rejected by the data ($\chi^2_3 = 12$). However, it turned out that the sum of the v_j was estimated to be almost exactly equal to the sum of the w_j (see row 3 of table 1), so this adding-up constraint was imposed in subsequent estimates.

The final generalization considered was to disaggregate personal income into its two main components—factor income and transfer payments:[12]

$$P_t = F_t + V_t. \tag{17}$$

This made the estimating equation:

$$\begin{aligned} C_t = k_0 &+ k_1 r_t Y_t + k_2(W_t - A_t) + \sum_{j=0}^{q} k_{3+j} A_{t-j} \\ &+ \sum_{j=0}^{n} w_j F_{t-j} + \sum_{j=0}^{n} \phi_j V_{t-j} - \sum_{j=0}^{n} v_j (T_{t-j} - \lambda X^j_t) \\ &+ \sum_{j=0}^{n} \beta_j \left[(1 - \lambda) X^j_t + \sum_{i=1}^{m} (1 - D^i_t) S^i_{t-j} \right] + u_t. \end{aligned} \tag{18}$$

Once again, the null hypothesis that the ϕ_j (spending coefficients for transfers) are in fact equal to the w_j (spending coefficients for factor income) was tested by a χ^2 test. And once again it was resoundingly rejected ($\chi^2_3 = 19$; see row 4 of table 1). This time, however, the data also rejected the adding-up constraint $\Sigma \phi_j = \Sigma w_j$, which therefore was not imposed (table 1, row 5). In fact, most of the difference between the ϕ_j and the w_j was in their sums; the time patterns were remarkably similar.

To summarize these tests, we are left with a model that assigns distributed lag weights w_j to factor income, ϕ_j to transfers, v_j to regular taxes, and β_j to windfalls. The Σv_j and Σw_j are apparently equal, but the other sums are not.

V. Issues in Estimation

1. *Data*

Following the suggestion of Darby (1975), I used consumer expenditures, rather than pure consumption, as C_t. This seems most appropriate where the focus is on the evaluation of stabilization policy, as it is here, rather than on testing the PIH. Furthermore, it

[12] For this purpose, both employer contributions and business transfers are considered to be factor income, and the aspects of the 1975–76 tax cuts that are classified as transfer payments in the national income accounts are grouped with temporary taxes.

TABLE 2

EFFECTS ON DISPOSABLE INCOME OF 1975–76 TAX CUTS AND TRANSFERS*

	TAXES†		TRANSFERS‡		
QUARTER	Tax Rate Cuts	Rebate	Social Security	Earned Income	TOTAL
1975:2	8.5	31.2	6.7	0	46.4
1975:3	12.3	0	0	0	12.3
1975:4	11.9	0	0	0	11.9
1976:1	14.2	0	0	1.9	16.1
1976:2	14.0	0	0	1.6	15.6

* In billions of current dollars, at annual rates.
† From U.S. Bureau of Economic Analysis (February 1976, March 1977).
‡ Kindly supplied to the author by Joseph C. Wakefield, of the Bureau of Economic Analysis, in conversation.

avoids many complicated issues of definition (e.g., Which goods are durables? How fast do they depreciate? etc., etc.). The cost of this shortcut is that the theoretical interpretation of some of the parameters is lost. For example, k_1 includes the effects of r_t on both pure consumption (which may be positive or negative) and spending on durables (which should be negative). Similarly, the lag weights w_j might be expected to be less smooth than the lag of pure consumption behind income because of the lumpy nature of durables.[13]

Data for the 1968 surcharge were taken from Okun (1971) and converted to 1972 dollars by the deflator for personal consumer expenditures; these comprise S_t^1. Data for the various 1975–76 tax cuts are shown in table 2; they were similarly deflated and segregated into two time series. The S_t^2 series is defined as the explicitly one-shot measures: the tax rebates and the social security bonuses (hereafter referred to as "the rebate"). The rest is considered as S_t^3. Since the 1975 cuts were extended several times and are now a permanent feature of the tax code, an arbitrary decision had to be made as to when they became "permanent."[14] I decided to cut off S_t^3 after 1976:2, because by then the cuts had already been extended once in the Revenue Adjustment Act of 1975 and a second time in the Tax Reform Act of 1976.

[13] To account for the special features of expenditures on consumer durables, several additional variables were tried in some earlier regressions. Neither the stock of durables, nor the relative price of durables, nor the unemployment rate succeeded in significantly lowering the sum of squared residuals. In several cases, the signs of the coefficients were even the opposite of what theory suggests. It may be that these variables are more relevant to the choice between saving in the form of durables vs. in financial form than they are to the choice between spending and saving.

[14] At an early stage of this research, I experimented with a learning model in which a temporary tax gradually came to be considered permanent as it remained on the books longer and longer. This experiment was unsuccessful.

Data on consumer net worth (W_t), and its breakdown into stock market (A_t) and non–stock market ($W_t - A_t$) components, were taken from the data bank of the MPS model and converted to 1972 dollars. They are based on a number of primary sources, the most important of which is the flow of funds.

Because of the recent findings of Boskin (1978), I thought it important to use a real after-tax interest rate. Since the construction of this series was a fairly involved affair, I explain it only in Appendix C.

2. *Distributed Lag Estimation*

The many distributed lags in equation (18) were estimated by an adaptation of the Almon (1965) lag technique, as a method of conserving on parameters. Generally, a third-degree polynomial with a zero constraint at the far end was used. Preliminary tests suggested that these end-point constraints ($w_{n+1} = 0$, etc.) could not be rejected. An unconstrained version of one specification, run as an experiment, showed that the polynomial constraint had very little effect on the w_j coefficients but did influence the β_j coefficients.

The distributed lag effects of assets were estimated as follows. In some preliminary regressions, the two components were combined and a cubic distributed lag over n quarters was estimated. Then the two components were disaggregated. These preliminary tests showed quite clearly (*a*) that the coefficients were not the same and (*b*) that the lag was much shorter than n quarters. (As explained just below, n was chosen to be 7.) In the case of non–stock market wealth, an estimated distributed lag over 4 quarters attached virtually all of the weight to the current (start of period) value, so all lagged values were omitted. In the case of stock market wealth, when a cubic was estimated over $j = 0, 1, \ldots, 7$, the coefficients turned out to be almost linear and to be virtually zero after $j = 2$ (a small positive value for $j = 3$ and small negative values for $j = 4, \ldots, 7$). Thus a linear distributed lag over $j = 0, \ldots, 3$ was adopted as the final specification.

The length of the distributed lag, n, was selected by running a preliminary version of the regression for alternative values of n ranging from 6 to 10. A very clear minimum in the sum of squared residuals was found around $n = 7$ or $n = 8$, with the former having a slightly better fit and slightly better coefficients. On this basis, $n = 7$ was selected for all subsequent work.

3. *Treatment of the 1975 Rebate*

As has been noted already, the model divides the 1975–76 tax cuts into two parts: S_t^2 includes the rebate, while S_t^3 includes all the rest.

This makes a strong (and questionable) assumption about how consumers treated the rebate. In particular, it assumes that they treated it just like the 1968 surcharge and the other 1975 reductions: essentially as if a fraction λ of it was a regular increase in income, while a fraction $1 - \lambda$ was a pure windfall. Given the nature of the rebate, this is questionable, to say the least.

An alternative assumption—equally strong as the first—is that consumers treated the rebate as a pure windfall right from the start.[15] While it is not obvious that this is true, since consumers might have anticipated a repeat performance with some reasonable probability, it does seem a plausible working hypothesis. Fortunately, it is not difficult to modify any of the three models to accommodate this alternative hypothesis; all that is necessary is that the lag weights β_j be applied to S_t^2 starting immediately in 1975:2. When this was done in one version of the model, the resulting equation had virtually an identical SSR, almost the same estimated λ, and very similar implications about spending patterns out of the rebate. Thus the conclusions of this study seem insensitive to the treatment of the rebate.

VI. Empirical Results

1. *Parameter Estimates and Interpretation*

Estimation was done by the numerical optimization package developed by S. M. Goldfeld and R. E. Quandt. The results from estimating equation (18) on quarterly data covering 1953:1–1977:4 are presented in table 3. The number in parentheses next to each estimated coefficient is its asymptotic standard error (or rather a numerical estimate thereof).

In interpreting the standard error of the regression, it should be mentioned that, as explained in Appendix B, the equation was actually transformed so that the left-hand variable was the average propensity to consume (APC), C_t/Y_t, rather than consumer spending. Thus, the standard error of .0038 is relative to a typical value for the APC of about .90. This represents an excellent fit.[16] At 1977 income levels, it translates to a standard error of about $3.5 billion in predicting consumer spending. Of course, obtaining a good fit with a consumption function is hardly a notable achievement, and the equation—like most consumption functions—does suffer from some autocorrelation ($\rho = .54$).

[15] Since the rebate was off the books by 1975:3, and hence treated as a windfall in any case, only 1975:2 is at issue here.

[16] The standard errors of a comparable equation in Modigliani and Steindel (1977) are .0056 with an autocorrelation correction and .0065 without.

TABLE 3

NONLINEAR CONSUMPTION FUNCTION PARAMETER ESTIMATES*

k_0	k_1	k_2	λ	ρ
30.1 (7.1)	.0002 (.0007)	.021 (.005)	.50 (.32)	.54 (.11)

DISTRIBUTED LAG COEFFICIENTS†

j	w_j	ϕ_j	v_j	β_j	k_{3+j}	γ_j	Rebate
0	.60 (.05)	.50 (.16)	.34 (.09)	−.03 (.26)	.009 (.002)	.16	.16
1	.16 (.02)	−.02 (.09)	.14 (.04)	−.03 (.11)	.006 (.001)	.06	−.03
2	−.06 (.03)	−.20 (.10)	.04 (.05)	−.01 (.09)	.004 (.001)	.02	−.01
3	−.12 (.02)	−.14 (.08)	.01 (.05)	.03 (.09)	.002 (.0005)	.02	.03
4	−.08 (.01)	.04 (.06)	.02 (.03)	.08 (.07)	...	.05	.08
5	.02 (.01)	.26 (.07)	.05 (.03)	.11 (.06)	...	.08	.11
6	.10 (.02)	.39 (.09)	.08 (.04)	.12 (.08)	...	.10	.12
7	.11 (.02)	.34 (.08)	.07 (.04)	.09 (.07)	...	.08	.09
Sum	.74	1.17	.74	.36	.021 (.005)	.55	.55

NOTE.—k_0 = constant, k_1 = coefficient of interest rate, k_2 = coefficient of non-stock market wealth, λ = weight on regular income, ρ = autocorrelation coefficient. Sum of squared residuals = .00147. SE = .00383, SE of unadjusted errors (without a correction for autocorrelation) = .00455, N observations = 100.

* Asymptotic SEs are in parentheses.

† Components may not add to totals due to rounding.

Turning to the coefficient estimates, the most critical parameter for purposes of this study is λ, the weight attached to regular income in equation (8). The point estimate of .50 suggests that temporary taxes that are still on the books are treated like 50-50 blends of windfalls and regular taxes. However, the standard error is regrettably large; there are, after all, pitifully few observations that can be used to estimate λ. The null hypotheses $\lambda = 0$ or $\lambda = 1$ can nonetheless be tested by likelihood ratio tests. When these tests were run with equation (18) as the unconstrained regression (see table 1, rows 6 and 7), the null hypothesis that $\lambda = 0$ (temporary taxes are regarded as pure windfalls) could not be rejected ($\chi_1^2 = 1.32$). But the null hypothesis that $\lambda = 1$ (temporary taxes are regarded as regular income) could be rejected if we were not too fussy about significance levels ($\chi_1^2 = 2.84$).

I turn next to the distributed lag coefficients of the various income terms. The w's for factor income are very large and positive at first, then turn small and negative, and finally become positive again at the end. This general shape accords well with our expectations.[17] The ϕ_j coefficients for transfer payments follow a similar shape but are much more erratic and less well pinned down econometrically. A notable

[17] Because of the estimating form, it is unlikely that simultaneity has much to do with the large estimate for w_0. See Appendix B.

feature is that their sum is nearly 1.2, indicating "overspending" during the first 2 years after receipt of a transfer payment.

The v_j coefficients for regular taxes also exhibit a characteristic U-shape, but in much more muted fashion. As compared with factor income, spending in the first year after a regular tax cut is apparently substantially less, after which it catches up. This was surprising at first, since Dolde (1976) and Modigliani and Steindel (1977) had suggested that regular tax changes are "more permanent" than regular income.[18] However, it turns out that the following second-order autoregressives describe the deviations from trend of personal income (P_t) and regular taxes (T_t):

$$P_t = \underset{(.09)}{1.32}P_{t-1} - \underset{(.09)}{.38}P_{t-2},$$

$$T_t = \underset{(.09)}{1.11}T_{t-1} - \underset{(.10)}{.21}T_{t-2}.$$

Using the formulas derived earlier for permanent income, these time-series models imply that "permanent personal income" and "permanent taxes" are given by $P_t^p = 14.9P_t - 5.7P_{t-1}$ and $T_t^p = 9.3T_t - 1.9T_{t-1}$, so that, contrary to Dolde and Modigliani and Steindel, we should actually expect a stronger short-run response of consumption to fluctuations in P_t than to fluctuations in T_t, which is exactly what I find.

The next column, the β_j's, are in some sense out-of-sample extrapolations since there are no "pure windfalls" recorded in the data.[19] Their only use is to form the weighted average $\lambda w_j + (1 - \lambda)\beta_j$, which is reported in the column marked "γ_j." These are the expenditure coefficients for income from a temporary tax that remains on the books for the entire 2-year horizon. To illustrate the opposite extreme, the column marked "Rebate" shows the spending coefficients for a temporary tax that lasts only 1 quarter. These two columns differ in details but are quite similar. There is a moderate spending response in the initial quarter, followed by very little spending over the next 3 or 4 quarters. Most of the spending out of a temporary tax cut, according to these estimates, comes 5 or more quarters after the cut.

The coefficient of assets (.02) is comparable to what others have estimated, though a bit on the low side. A likelihood ratio test of the null hypothesis $k_2 = 0$ (which, in this constrained form, also implies k_3

[18] But see Dolde (1979), where the transitory nature of allegedly permanent tax changes is stressed.

[19] I made an attempt, in some early regressions, to treat the National Service Life Insurance Dividends of 1950 in this way, but I was not successful.

$= \ldots = k_6 = 0$) produced a test statistic of 17.8, which is highly significant at any reasonable significance level (table 1, row 8).

If we ignore the fact that income from property is included in the measure of income, the parameter k_2 can be given an interesting theoretical interpretation. In the basic life-cycle model, the consumer maximizes a utility function of the form

$$\sum_{t=0}^{T} \left(\frac{1}{1+\rho}\right)^t \frac{C_t^{1-\delta}}{1-\delta},$$

subject to a lifetime wealth constraint. It can be shown by straightforward computations that the optimal solution for initial consumption is $C_0 = [(r - g)/(1 + r)]W$, where W is lifetime wealth and g is the optimal growth rate of C_t, defined as $g = [(1 + r)/(1 + \rho)]^{1/\delta} - 1$. This means that k_2 corresponds to the theoretical coefficient $(r - g)/(1 + r)$, which for small values of r and ρ is approximately equal to $r + (1/\delta)(\rho - r)$. Thus for small values of r the estimated value $k_2 = .02$ implies that the subjective discount rate, ρ, is approximately $.02\delta$ or around 2–3 percent per quarter for plausible values of δ.

One striking result, though it is peripheral to the subject of this study, is the tiny coefficient of the real after-tax interest rate.[20] This finding turned up in every specification of the model, including several alternative measures of the rate of interest. (Sometimes the coefficient was trivially negative, sometimes trivially positive, but always trivial.) While it accords well both with my earlier work (Blinder 1975) and with the work of others, it stands in sharp contrast with Boskin's (1978) recent finding of a strong positive interest elasticity of saving.

2. *Temporary versus Permanent Taxes*

We can now address the principal issue of this study: How effective are explicitly temporary income tax changes as compared to those announced to be permanent? Table 4 contains the answers derived from the model, using the parameter estimates presented in table 3 to make equation (4) operational and using an annual real interest rate of 3 percent in updating wealth according to equation (10). It can be seen from column 4 that a temporary tax is about one-half as effective as a permanent tax in the first year, rising to about three-quarters as effective in the second year. Spending out of a rebate is somewhat slower than this. My estimated cumulative spending propensities out

[20] The specific coefficient in table 3 means that a 1 percentage point rise in r lowers savings by about .02 of 1 percent of disposable income—a trivial amount.

TABLE 4

RELATIVE EFFECTIVENESS OF TEMPORARY TAXES

	CUMULATIVE SPENDING PROPENSITIES			RATIOS	
j	Permanent (1)	2-Year (2)	Rebate (3)	(2)/(1) (4)	(3)/(1) (5)
0	.34	.16	.16	.47	.47
1	.50	.23	.14	.46	.28
2	.55	.26	.16	.47	.29
3	.56	.30	.21	.54	.38
4	.59	.36	.30	.61	.51
5	.65	.46	.43	.71	.66
6	.73	.57	.56	.78	.77
7	.81	.65	.66	.80	.81

of a rebate are larger than those estimated by Modigliani and Steindel (1977) for the first few quarters but smaller thereafter.

These findings carry two important messages to fiscal policy planners. First, and most obvious, is that temporary taxes are less powerful devices for short-run stabilization purposes than are permanent ones. Second, and perhaps almost as important, the short-run relative ineffectiveness of such taxes implies that the impact of these measures in the second year is larger than might be expected. For example, according to table 4, each $1.00 of permanent tax reduction adds $0.25 to spending in the second year, while each $1.00 of a rebate adds $0.45. If the need is for a truly short-run stimulus to aggregate demand, this effect may also be unwanted.

Both of these points can be illustrated by examining what the equations have to say about the 1975–76 episode. First, it is useful to display the observed APCs for this period in table 5. There are two obvious phenomena crying out for explanation in these data. First, why did the APC drop so sharply in 1975:2? Second, why did it thereafter begin a steady climb to what is a truly extraordinary level by 1977:1? (The corresponding personal savings rate was only 4.2

TABLE 5

AVERAGE PROPENSITIES TO CONSUME, 1975–77

	1975	1976	1977
First quarter	.913	.914	.935
Second quarter	.881	.918	.926
Third quarter	.903	.922	.922
Fourth quarter	.907	.926	.925

SOURCE.—U.S. Bureau of Economic Analysis (various issues).

TABLE 6

ESTIMATED EFFECTS OF THE 1975–76 TEMPORARY TAX CUTS ON CONSUMER EXPENDITURES*

Quarter	Estimated Spending Effect
1975:2	5.9
1975:3	1.6
1975:4	2.7
1976:1	4.8
1976:2	6.7
1976:3	6.7
1976:4	7.1
1977:1	6.6
1977:2	7.2
1977:3	7.5
1977:4	7.3

* In billions of 1972 dollars.

percent—the lowest figure that had then been recorded since the Korean War.)

According to the estimates presented in this paper, the temporary tax cuts of 1975–76 contributed to both phenomena. Using the spending coefficients presented in table 4, table 6 shows the estimated direct (excluding multiplier) effects on consumer spending of the tax cuts of 1975:2 through 1976:2, inclusive.[21] It appears that (1) very little of the rebate was spent in 1975:2, (2) rather little of the disposable income attributable to the temporary tax cut package was spent during the remainder of 1975, and (3) more spending out of the temporary tax cuts was done in 1976 and yet more in 1977. Both the low APC of 1975:2 and the high APCs of late 1976 and early 1977 are tracked very well by the model. One observation which virtually jumps from table 6 is how very small these estimated spending impacts are relative to the size of the economy they were meant to stimulate (real GNP in the neighborhood of $1,250 billion).

Finally, there is one more question. If, instead of the 1975:2–1976:2 package of temporary measures, the government had cut taxes "permanently" in 1975:2 and then restored them to their original level starting in 1976:3, how large a tax cut would have achieved the same average effect on aggregate demand?[22]

Table 7 summarizes the model's answers to this question for three different choices of the horizon over which the "average effect on aggregate demand" might be defined. The first column gives the average direct impact on spending attributed by the model to the

[21] The reader is reminded that only these 5 quarters are considered temporary cuts.

[22] For this calculation I assume that consumers were successfully fooled into thinking that the 5-quarter tax cut would be permanent.

TABLE 7

EQUIVALENT PERMANENT TAXES*,†

Horizon	Average Impact on Spending of Actual 1975–76 Cuts	Cumulative Revenue Loss over 5 Quarters of Permanent Tax Cut with Equal Average Impact on Spending
4 quarters	3.7	9.5
6 quarters	4.7	12.4
8 quarters	5.3	15.9

* See text for definition.
† In billions of 1972 dollars, at annual rates.

1975–76 tax cuts. The next column shows how much total tax revenue the government would have had to relinquish during the same 5-quarter period (1975:2–1976:2) in order to achieve the same direct impact on spending through a permanent tax cut. Since the total 5-quarter revenue loss from the 1975–76 package was $20 billion, these numbers mean, for example, that a permanent tax cut about half as large ($9.5 billion vs. $20 billion) would have had the same first-year effect on aggregate demand. Over a 2-year horizon, however, the 1975–76 package had about 80 percent as much "bang for the buck" as a permanent tax cut.

VII. Summary

Both economic theory and casual empirical observation of the U.S. economy suggest that short-run spending propensities from temporary tax changes are smaller than those from permanent ones, but neither provides much guidance about the magnitude of this difference. This paper offers new empirical estimates of this difference and finds it to be quite substantial.

The analysis is based on an amendment of the standard distributed lag version of the PIH that distinguishes temporary taxes from other income on the grounds that the latter is "more transitory." This amendment, which is broadly consistent with rational expectations, leads to a nonlinear consumption function.

Though the standard error is unavoidably large, the point estimate suggests that a temporary tax change is treated as a 50-50 blend of a normal income tax change and a pure windfall. Over a 1-year planning horizon, a temporary tax change is estimated to have only a little more than half as much impact as a permanent tax change of equal magnitude, and a rebate is estimated to have only about 38 percent as much impact. The model tracks both the extraordinarily high savings rate of 1975:2 and the extraordinarily low savings rates of late 1976 and early 1977 very well and attributes part of both phenomena to the temporary tax measures of 1975–76. Finally, it is estimated that a permanent tax cut of about \$9.5 billion (in 1972 dollars) would have had the same impact on aggregate demand over the first 4 quarters as the \$20 billion of 1975–76 tax cuts.

Appendix A

The Lifetime Budget Constraint

This Appendix demonstrates a result that, to my knowledge, is not very well known: that the long-run MPC corresponding to a consumption function of the form

$$C_t = kA_t + \sum_{j=0}^{n} b_j Y_{t-j} \tag{A1}$$

is unity for any value of k (greater than the real rate of interest) and for any b's.

Proof: Write equation (10) in the text as

$$[1 - (1 + r)L]A_t = Y_{t-1} - C_{t-1}, \tag{A2}$$

and write (A1) as

$$C_t = kA_t + b(L)Y_t, \tag{A3}$$

where L is the lag operator and $b(L) = b_0 + b_1 L + \ldots + b_n L^n$. Applying the operator $1 - (1 + r)L$ to (A3) and using (A2) yields $[1 - (1 + r - k)L]C_t = \{kL + [1 - (1 + r)L]b(L)\}Y_t$, which can be written

$$C_t = B(L)Y_t, \tag{A4}$$

where

$$B(L) = \frac{kL + [1 - (1 + r)L]b(L)}{1 - (1 + r - k)L}. \tag{A5}$$

To obtain the lifetime spending generated by a \$1.00 impulse in Y_t, we must compute the discounted sum of coefficients:

$$\text{MPC} = \sum_{i=0}^{\infty} \frac{B_i}{(1 + r)^i}. \tag{A6}$$

To simplify the notation, let $\theta \equiv 1 + r - k$. Assuming that $\theta < 1$ (i.e., that $k > r$), (A5) can be written: $B(L) = \{kL + [1 - (1 + r)L]b(L)\}(1 + \theta L + \theta^2 L^2 + \ldots)$. This is of the form $B(L) = B_0 + B_1 L + B_2 L^2 + \ldots$, with the following coefficients:

$$
\begin{aligned}
B_0 &= b_0, \\
B_1 &= \theta b_0 + [b_1 + k - (1 + r)b_0], \\
B_2 &= \theta^2 b_0 + \theta[b_1 + k - (1 + r)b_0] + [b_2 - (1 + r)b_1], \\
&\ \ \vdots \\
B_n &= \theta^n b_0 + \theta^{n-1}[b_1 + k - (1 + r)b_0] + \ldots + [b_n - (1 + r)b_{n-1}], \\
B_{n+1} &= \theta^{n+1} b_0 + \theta^n[b_1 + k - (1 + r)b_0] + \ldots + \theta[b_n - (1 + r)b_{n-1}] \\
&\quad - (1 + r)b_n, \\
B_{n+1+s} &= \theta^s B_{n+1}, \qquad s = 1, 2, \ldots .
\end{aligned}
$$

Substitution of all of these into (A6), and some truly horrendous grinding, establishes that MPC = 1 regardless of the magnitudes of k and the b_j (as long as $k > 0$). Q.E.D.

A brief word on the interpretation of the sum of the b_j in (A1) may be in order here. Should Y_t rise permanently by \$1.00—a statement that is basically meaningless if the autoregressive process assumed in the text (eq. [2]) really holds—the eventual change in C_t would, by (A4), be $B(1) = \Sigma_{i=0}^{\infty} B_i$. According to (A5), this sum is

$$B(1) = \frac{k - r\sum_{j=0} b_j}{k - r} = 1 + \frac{r}{k - r}\left(1 - \sum_{j=0}^{n} b_j\right).$$

Thus Σb_j controls the size of the spending response to a hypothetical permanent rise in income; it does not influence the lifetime MPC.

Appendix B

Details on the Estimating Equation

This Appendix derives and explains the equation that was actually estimated. I begin by repeating equation (14) of the text:

$$C_t = k_0 + k_1 r_t Y_t + k_2(W_t - A_t) + \sum_{j=0}^{q} k_{3+j} A_{t-j}$$
$$+ \sum_{j=0}^{n} w_j(Y_{t-j} + \lambda X_t^j - S_{t-j}) + \sum_{j=0}^{n} \beta_j \left[(1 - \lambda)X_t^j + \sum_{i=1}^{m} (1 - D_t^i)S_{t-j}^i\right] + u_t.$$

For purposes of reducing heteroscedasticity, the assumption was made that the standard deviation of u_t was proportional to Y_t, so the whole equation was divided through by Y_t to get

$$\begin{aligned}\mathrm{APC}_t = {} & \frac{k_0}{Y_t} + k_1 r_t + k_2\left(\frac{W_t}{Y_t} - \frac{A_t}{Y_t}\right) + \sum_{j=0}^{q} k_{3+j} \frac{A_{t-j}}{Y_t} \\ & + \sum_{j=0}^{n} w_j(z_{t-j} - q_{t-j}) + \sum_{j=0}^{n} \beta_j \left[(1 - \lambda)x_t^j + \sum_{i=1}^{m} (1 - D_t^i)s_{t-j}^i\right] + \epsilon_t,\end{aligned} \tag{B1}$$

where

$$z_{t-j} \equiv Y_{t-j}/Y_t \quad (\text{note: } z_t = 1 \text{ for all } t),$$

$$x_t^j = X_t^j/Y_t,$$

$$s_{t-j} = \frac{S_{t-j}}{Y_t},$$

$$q_{t-j} = s_{t-j} - \lambda x_t^j,$$

$$\epsilon_t = \frac{u_t}{Y_t}.$$

To estimate (B1), the assumption was made that both w_j and β_j follow third-degree polynomials in j:

$$w_j = a_0 + a_1 j + a_2 j^2 + a_3 j^3, \qquad \text{(B2)}$$
$$\beta_j = b_0 + b_1 j + b_2 j^2 + b_3 j^3.$$

The end-point constraints mentioned in the text ($w_{n+1} = \beta_{n+1} = 0$) are thus

$$a_0 + (n+1)a_1 + (n+1)^2 a_2 + (n+1)^3 a_3 = 0, \qquad \text{(B3)}$$
$$b_0 + (n+1)b_1 + (n+1)^2 b_2 + (n+1)^3 b_3 = 0.$$

Equations (B3) were used to eliminate the parameters a_0 and b_0. The adding-up constraint discussed (and rejected) in the text was thus

$$\sum_{j=0}^{n} \{a_1[j-(n+1)] + a_2[j^2-(n+1)^2] + a_3[j^3-(n+1)^3]\},$$
$$= \sum_{j=0}^{n} \{b_1[j-(n+1)] + b_2[j^2-(n+1)^2] + b_3[j^3-(n+1)^3]\}, \qquad \text{(B4)}$$

which was used to eliminate the parameter a_1.

Finally, in estimating (B1), ϵ_t was assumed to follow a first-order autoregressive scheme, $\epsilon_t = \rho\epsilon_{t-1} + e_t$, where e_t is white noise. Estimation was by nonlinear least squares, which is equivalent to maximum likelihood if e_t is normally distributed. The function actually minimized was

$$\sum_{t=1}^{T} \Bigg\{ \mathrm{APC}_t - \frac{k_0}{Y_t} - k_1 r_t - k_2\left(\frac{W_t - A_t}{Y_t}\right) - \sum_{j=0}^{q} k_{3+j}\frac{A_{t-j}}{Y_t} - \sum_{j=0}^{n} w_j(z_{t-j} - q_{t-j}) - \sum_{j=0}^{n} \beta_j\left[(1-\lambda)x_t^j - \sum_{i=1}^{m}(1-D_t^i)s_{t-j}^i\right] - \rho\epsilon_{t-1}\Bigg\}^2,$$

with all the above-mentioned definitions and parameter restrictions substituted in.

The estimating forms when regular income was further disaggregated were derived in precisely analogous ways from equations (16) and (18).

Appendix C

Calculation of the Real After-Tax Interest Rate

The real after-tax interest rate is defined as $r_t = i_t(1-\tau_t) - \pi_t$, where i_t is the nominal interest rate, τ_t is the marginal tax rate, and π_t is the expected rate of inflation.

Nominal interest rate.—Four different nominal interest rates were tried: a corporate bond rate, the 4–6 month commercial paper rate, the 3-month Treasury bill rate, and a weighted average of rates received on various time and savings accounts. While all four led to very similar results, the Treasury bill rate gave the best fit and so was adopted.

Marginal tax rate.—The marginal tax rate was created from the average tax rate in the following way. Let $T(y)$ be the tax function facing an individual taxpayer, and let $f(y)$ be the frequency distribution of income. The only directly observable tax rate is the aggregate average rate, which is $A = [\int T(y)f(y)dy]/[\int yf(y)dy]$.

Now suppose the whole income distribution shifts to the right with no change in its shape; that is, it shifts from $f(y)$ to $f[y(1-h)]$, where h connotes a "small" multiplicative shift. The average income is then $Y(h) = \int yf[y(1-h)]dy$, so that, for small shifts (h near zero), $dY/dh = -\int y^2 f'(y)dy$. Similarly, average tax payments after the shift are $T(h) = \int T(y)f[y(1-h)]dy$, so that $dT/dh = -\int yT(y)f'(y)dy$. The aggregate marginal tax rate, M, is the ratio of these: $M \equiv [\int yT(y)f'(y)dy]/[\int y^2 f'(y)dy]$.

To evaluate these integrals and obtain a closed expression for M/A, I adopted the following functional forms:

$$T(y) = ay^b \qquad (b > 1, a > 0),$$
$$f(y) = \gamma e^{-\gamma y} \qquad (\gamma > 0).$$

With these assumptions, the two ratios of integrals work out to be $A = (a\gamma^2/\gamma^{b+1})\Gamma(b+1)$, $M = (a\gamma^3/2\gamma^{b+2})\Gamma(b+2)$, where $\Gamma(n)$ is the "gamma function," namely, $\Gamma(n+1) = n\Gamma(n)$. Hence $M/A = (b+1)/2$, and using a tax elasticity of $b = 1.6$ gives $M = (1.3)A$.

The average tax rate was computed, quarter by quarter, by dividing the sum of federal and state-local personal tax and nontax payments by personal income excluding transfers (an approximation to the tax base). These are all official national income accounts series.

Expected rate of inflation.—Inflationary expectations were generated by a model based on what has been called "economically rational" expectations.[23] The idea is that agents, in informing themselves, begin with the data that are cheapest per unit of informational content and then proceed to process more costly data until the marginal cost and (expected) marginal benefits are equated. In this particular application, I assumed that consumers base their expectations of the inflation rate ($\dot{P}_t$) on its own past history and on the history of the growth rate of the money supply ($\dot{M}_t$). Thus I estimated an equation

$$\dot{P}_t = a_0 + \sum_{j=1}^{J} a_j \dot{P}_{t-j} + \sum_{i=0}^{I} b_i \dot{M}_{t-i} + e_t$$

on actual U.S. quarterly data, using the deflator for personal consumption expenditures for P_t and M_2 for M_t, and assumed that consumers used this equation to generate expectations. In estimation, I used the Almon (1965) lag technique with third-degree polynomials, no end-point constraints, and various choices for J and I. The best results were obtained with $J = 11$ and $I = 17$, namely[24]

$$\underset{(.38)}{\dot{P}_t = -.60} + \sum_{1}^{11} a_j \dot{P}_{t-j} + \sum_{0}^{17} b_i \dot{M}_{t-i}, \; R^2 = .73, \qquad \text{D-W} = 1.98,$$

standard error = 1.42, mean of dependent variable = 3.35,

$$\underset{(.10)}{\Sigma a_j = .67} \qquad \underset{(.09)}{\Sigma b_i = .27.}$$

[23] See Feige and Pearce (1976).

[24] Standard errors are in parentheses. The D-W is the Durbin-Watson statistic. The period of estimation was 1951:3–1977:4, the longest period possible given the need for 17 lagged values of M.

References

Almon, Shirley. "The Distributed Lag between Capital Appropriations and Expenditures." *Econometrica* 33 (January 1965): 178–96.

Blinder, Alan S. "Distribution Effects and the Aggregate Consumption Function." *J.P.E.* 83, no. 3 (June 1975): 447–75.

———. "Intergenerational Transfers and Life Cycle Consumption." *A.E.R. Papers and Proc.* 66 (May 1976): 87–93.

Blinder, Alan S., and Solow, Robert M. "Analytical Foundations of Fiscal Policy." In *The Economics of Public Finance*, by Alan S. Blinder and others. Washington: Brookings Inst., 1974.

Boskin, Michael J. "Taxation, Saving, and the Rate of Interest." *J.P.E.* 86, no. 2, pt. 2 (April 1978): S3–S27.

Darby, Michael R. "The Allocation of Transitory Income among Consumers' Assets." *A.E.R.* 62 (December 1972): 928–41.

———. "Postwar U.S. Consumption, Consumer Expenditures, and Saving." *A.E.R. Papers and Proc.* 65 (May 1975): 217–22.

Dolde, Walter. "Forecasting the Consumption Effects of Stabilization Policies." *Internat. Econ. Rev.* 17 (June 1976): 431–46.

———. "Capital Markets and the Short Run Behavior of Life Cycle Savers." *J. Finance* 33 (May 1978): 413–28.

———. "Temporary Taxes as Macro-economic Stabilizers." *A.E.R. Papers and Proc.* 69 (May 1979): 81–85.

Eisner, Robert. "Fiscal and Monetary Policy Reconsidered." *A.E.R.* 59 (December 1969): 897–905.

Feige, Edgar L., and Pearce, Douglas K. "Economically Rational Expectations: Are Innovations in the Rate of Inflation Independent of Innovations in Measures of Monetary and Fiscal Policy?" *J.P.E.* 84, no. 3 (June 1976): 499–522.

Foley, Duncan K., and Hellwig, Martin F. "Asset Management with Trading Uncertainty." *Rev. Econ. Studies* 42 (July 1975): 327–46.

Friedman, Milton. *A Theory of the Consumption Function.* Princeton, N.J.: Princeton Univ. Press (for Nat. Bur. Econ. Res.), 1957.

Goldfeld, Stephen M., and Quandt, Richard E. *Nonlinear Methods in Econometrics.* Amsterdam: North-Holland, 1972.

Hall, Robert E. "Stochastic Implications of the Life Cycle–Permanent Income Hypothesis: Theory and Evidence." *J.P.E.* 86, no. 6 (December 1978): 971–87.

Houthakker, Hendrik S., and Taylor, Lester D. *Consumer Demand in the United States.* Cambridge, Mass.: Harvard Univ. Press, 1966.

Juster, F. Thomas. "A Note on Prospective 1977 Tax-Cuts and Consumer Spending." Unpublished paper, Univ. Michigan, January 1977.

Lucas, Robert E., Jr. "Econometric Policy Evaluation: A Critique." In *The Phillips Curve and Labor Markets*, edited by Karl Brunner and Allan H. Meltzer. Amsterdam: North-Holland, 1976.

Modigliani, Franco, and Brumberg, Richard E. "Utility Analysis and the Consumption Function: An Interpretation of Cross-Section Data." In *Post-Keynesian Economics*, edited by Kenneth K. Kurihara. New Brunswick, N.J.: Rutgers Univ. Press, 1954.

Modigliani, Franco, and Steindel, Charles. "Is a Tax Rebate an Effective Tool for Stabilization Policy?" *Brookings Papers Econ. Activity*, no. 1 (1977), pp. 175–203.

Muth, John F. "Optimal Properties of Exponentially Weighted Forecasts." *J. American Statis. Assoc.* 55 (June 1960): 299–306.

Okun, Arthur M. "The Personal Tax Surcharge and Consumer Demand, 1968–70." *Brookings Papers Econ. Activity*, no. 1 (1971), pp. 167–211.

Sargent, Thomas J. "Rational Expectations, Econometric Exogeneity, and Consumption." *J.P.E.* 86, no. 4 (August 1978): 673–700.

Springer, William L. "Did the 1968 Surcharge Really Work?" *A.E.R.* 65 (September 1975): 644–59.

U.S. Bureau of Economic Analysis. *Survey of Current Business.* Washington: Government Printing Office, various issues.