Short-Run Asset Effects on Household Saving and Consumption

by

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I. Introduction and Summary

This paper introduces a new approach to the measurement of the short-run effects of changes in the value of assets on household saving and consumption. Previously, the timing and magnitude of asset effects on consumer behavior have been estimated exclusively from time-series data--either from a single equation consumption function or, more commonly in recent years, from a large-scale econometric model. The basic difficulty with these econometric models based on aggregate time-series data is the small number of independent observations available for determining the appropriate parameters in a model with a large number of structural equations and a large number of alternative forms to select among for each equation. With the use of increasingly complex lag structures, the number of alternative forms to choose from for each structural equation has proliferated greatly. Small variations in the forms fitted will frequently have very little effect on the goodness of fit of an equation in the model but can change dramatically the distributed lag pattern.

Survey data, covering a large cross-section of consumers (or business organizations if attention is focused on investment rather than consumption behavior), provide a basis for correcting this deficiency. This does not mean that such data do not have deficiencies of their own, but simply that the paucity of independent observations characterizing time-series, compounded by temporal changes in the economic structure, cannot be remedied without the use of cross-section information. The most serious deficiencies in survey or other cross-section data can be corrected with the availability of sufficient resources. Continuous cross-section or panel data covering the same households (or groups of households) over several consecutive periods of time are necessary to avoid some of the limitations of a single cross-section, but highly useful information on short-run asset effects on saving and consumption can be obtained from appropriate information collected in a single cross-section.

The cross-section data used for this paper consist of information on saving, income, and assets for a sample of households in 1963 (with income data for 1962 as well as 1963) collected by the Federal Reserve Board (FRB) in their surveys of the Financial Characteristics of Consumers and Changes in Family Finances.[1] Saving in these surveys is measured from data on changes in assets and liabilities (apart from revaluation items) but is, of course, conceptually equal to income minus consumption expenditures. The surveys, which oversampled upper income groups, collected detailed information for 2,159 households in all income classes not only for income, saving and the values of major categories of assets held at the beginning and end of the year but also for the amounts of individual stocks held at the beginning of the year. For purposes of this paper, we were able to use the data on individual stockholdings to estimate capital gains (largely unrealized) on a major, and the most volatile, part of household wealth. As a result, the effect on saving of capital gains during the year can be studied on the basis of a large number rather than a handful of observations.

The economic importance of obtaining further insights into the short-run impact of capital gains on saving is highlighted by the

MIT-Pennsylvania-SSRC (MPS) model, which is in our opinion the best of the existing large-scale models in the United States for assessing the impact of monetary policy.[2] In that model, one of the most important channels of transmission of monetary policy is the wealth effect on consumption. Monetary policy affects short-term and long-term interest rates, dividend yields and hence the value of assets, and--of particular importance in view of its magnitude and volatility--the value of common stock holdings. Qualitatively, there can be little question that assets affect consumption, and on the assumption of a life-cycle saving theory it is possible to apply some rough checks of reasonableness to the MPS estimate of the long-run effect on consumption of a change in the value of assets. However, the magnitude and, even more so, the time-sequence of this effect is extremely difficult to determine from time-series data.

The updated (January 1973) version of the MPS model implies that an increase of one dollar in stock value increases the long-run rate of consumption by $.054, with most of this effect taking place in the first year. However, while the estimated long-run effect has changed only moderately in different recent versions of the model, the timing of that effect has changed greatly. Moreover, as discussed subsequently, other time-series analyses have provided radically different results for the magnitude and timing of asset effects on consumption. Given the very large margin of indeterminancy in estimating the pattern of asset effects on consumption from time-series data, the survey results should as a minimum serve as a highly valuable check on the time-series analysis.

Our analysis of the survey data indicates that household saving in 1963 was reduced by about $.02 to $.04 for every dollar of capital gains on corporate stocks during that year, with the precise figure obtained depending on the form of the regression utilized and the nature of the adjustments made to the original FRB sample. The midpoint of this range, \$.03, is considerably below the \$.040 figure implied by the latest published version of the MPS model but is only modestly less than the \$.032 implied by the updated (January 1973) and as yet unpublished version of the model. The midpoint of the survey estimates is, however, well above the corresponding figures implied by other time-series analyses by Arena (1964,1965), Bhatia, Feldstein and Feldstein and Fane.

While estimates of long-run asset effects on household saving are also provided by our analysis of the 1963 survey data, they are not as stable as the short-run asset (or capital gains) coefficients. Moreover, as is well known, the long-run asset coefficients in saving regressions based on household survey data are downward biased in absolute value since assets serve as a proxy for savings tastes as well as a measure of resources available for consumption and investment. As a consequence, in the absence of panel information, cross-section data are not very useful in casting additional light on the wide assortment of estimated long-run asset effects obtained in different time-series analyses.

II. Earlier Work

The only extensive use of cross-section data in earlier studies of asset effects on household savings and consumption consisted of the estimation of long-run asset effects. These effects were generally rather small, even when they had the theoretically expected size and were statistically significant.[3] They are understated because of differences in household saving tastes which cannot adequately be kept constant in an analysis of single cross-section data.

Earlier time-series studies of asset effects included estimates both of long-run and short-run effects. Ando and Modigliani (1963,1964) regressed aggregate annual U.S. consumption on disposable labor income and beginning of year personal assets or wealth in a single equation model covering the period 1929-56 exclusive of 1941-46, and in different forms of the equation tested obtained asset coefficients ranging from .04 to .08. These compared with "theoretically expected" coefficients ranging from .11 to .13 on the basis of the life-cycle hypothesis (if yields on assets vary between three precent and five percent and the annual rate of growth of aggregate income between zero and four percent). If the consumption function is assumed homogeneous in income and assets, most of the asset coefficients ranged from about .06 to .08, and Ando and Modigliani concluded that the most likely values were between .072 and .08. No distinction was made in these regressions between long-run and short-run asset effects.

Arena subsequently included contemporaneous stock market gains as well as beginning of year personal assets or wealth in similar regressions covering the 1946-58 period on an annual basis and the 1947-63 period on a quarterly and annual basis. For the 1946-58 period, Arena obtained wealth coefficients ranging from .035 to .093 and capital gains coefficients from .013 to .036, though the latter were in general not statistically significant. For the 1947-63 period, the wealth coefficients fell in the same range (in those regressions where it was included) but the capital gains coefficients were much smaller (sometimes negative) and completely insignificant.

Evan's work cast some doubt even on the long-run asset effect on aggregate consumption. An annual time-series regression for the 1929-62 period exclusive of 1942-46 implied a highly significant .057 coefficient for initial welath, but a quarterly analysis for the post-World War II period pointed to an extremely small and completely insignificant .003 coefficient. However, most subsequent work, including quarterly as well as annual time-series analysis, has confirmed both a long-run asset as well as a capital gains effect on consumption.

Bhatia, Feldstein and Feldstein and Fane, who respectively derived aggregate consumption functions from annual U.S. time-series data for the period 1948-64, from annual U.S. time-series for 1929-66 excluding 1942-46, and from annual U.K. time-series data for 1948-69, obtained significant long-run asset effects. However, the relevant coefficients varied widely among the three analyses, ranging between .050 to .062 in the first analysis, .010 to .024 in the second, and .016 to .017 in the third. Feldstein and Fane, who included the current year's accrued (largely unrealized) capital gains on all assets as a separate variable, obtained a significant capital gain coefficient of .019. Bhatia's same year unrealized capital gain coefficient was somewhat smaller and Feldstein's same year accrued capital gain coefficient was insignificantly different from zero.

The consumption function in the MPS model represents a considerable refinement of the earlier work by Ando and Modigliani. It is fitted to quarterly data, now breaks down assets into stocks and all other assets and makes use of distributed lags in the asset variables. The updated version of the model implies a long-run stock coefficient of .054,[4] taking eight quarters to achieve, with 30.2 percent of the total impact achieved in the first quarter, 53.6 percent in the first half and 83.6 percent in the first year. The long-run stock coefficient is constrained to be equal to the other assets coefficient but no lag is allowed for in the other assets variable in which a change is assumed to achieve its ultimate effect on consumption in the same quarter. It is noteworthy that the long-run effect of stock assets on consumption in the most recent version of the MPS model is lower than in the preferred versions of the earlier Ando-Modigliani equations which in turn were lower than the values implied by their theory. It is also interesting that the time pattern of the total impact of changes in stock and other assets is quite different in the most recent version of the MPS model compared with earlier versions of that model, with a considerably slower effect on changes in the value of stocks held and a much faster impact of changes in the value of other assets.

Perhaps enough has been said to indicate the large margin of error in estimating the time-pattern of asset effects on consumption from time-series data even though asset effects are generally believed to play a critical role in the effectuation of monetary policy. To obtain reasonably definitive insights into the timing and magnitude of the wealth effect on consumption, it will probably be necessary to collect new data through surveys which compile continuous cross-section information on household savings, income, assets and capital gains. However, even with the cross-section information already available, it is possible to improve on our present knowledge of asset effects on consumption.

III. Limitations of New Data

There are several difficulties in extending to the corresponding time-series relationships the result of an analysis of survey data on personal saving, income, capital gains and assets derived from a single cross-section of households. First, since transitory components of some of these variables are likely to be relatively more important in cross-section data, there is the problem of separating out normal from transitory components, and this will be attempted in the subsequent analysis. Second, since we are interested in drawing intertemporal conclusions from interpersonal data, tastes effects must be considered. Third, measurement errors are generally larger for household cross-section than for aggregate time-series data, reflected both in larger random errors and in substantial survey under-estimation of savings.[5] However, there is reason to believe that the survey under-estimation of savings is not likely to affect seriously the shape of the saving-income relationship.[6] Moreover, for the value of stock assets held by households, and the associated capital gains, the FRB surveys are likely to be more accurate than the time-series data in the U.S.--at least for the high proportion of respondents who reported the number of shares owned in each stock in their portfolios.[7] Fourth, the definition of saving used in these surveys differs from the time-series (and customary economic) definition in several respects, of which the two most important are the absence of any deduction for depreciation on owned homes and of any allowance for saving in the form of life insurance. In our statistical results presented in the following section, we have consistently subtracted depreciation from the FRB measure of saving, and in part of our analysis we have added back a rough allowance for was restricted to those households which either reported no shareholdings or for which identified shareholdings accounted for at least 90 percent of the total value of their reported stockholdings. This restriction reduced the sample by 389 to 1,770 observations, and for part of the analysis other restrictions were also employed reducing this sample further. When stock was unidentified, capital gains were estimated at 20 percent for the year.[11] As a result, the capital gains variable contains an additional small random error which would tend to bias somewhat its estimated coefficient towards zero. The sensitivity of the capital gains coefficient to random errors is tested in our statistical analysis.

Finally, the weighting and reweighting procedures followed are important in the interpretation of the results because the survey heavily oversampled upper income groups and also because non-random subsamples are used in most of our analysis. Roughly 128 weight classes distinguished by income, age, retirement status and moving status are employed with each observation assigned a weight to represent a specific number of households in the population. Whenever observations are deleted, the remaining observations are proportionately scaled upward so that the total weight for each class remains unaffected. Should a weight class lose all of its observations, its weight is distributed among the members of a nearby weight class.

IV. New Statistical Results

A number of cross-section regressions were estimated from the FRB survey data using the general form $S = a + bY_N + cY_T + dZ + eA_{-1} + fAge + g\,Fam\,Size + hOccu$, where S is saving for 1963 net of depreciation saving in the form of life insurance.[8]

Since the major contribution of this paper is to make use of the unique FRB data on the detailed stock holdings of individual households for deriving new estimates of capital gains effects on saving, it is important to indicate that the capital gains data though of relatively high quality are still subject to some error. Capital gains are estimated from the prices at the beginning and end of the year for individual stocks held by each household at the beginning of the year. Even where respondents correctly reported the number of shares owned of each stock in their portfolios, our estimates of accrued capital gains (or losses) attributes to a household capital gains which accrued to a stock after it had been sold, while accrued capital gains on purchases made during the year are excluded. Even so, the error introduced is rather small. As measured by the FRB survey, realized capital gains on securities (which includes some gains on assets other than stocks) represent only about five percent of the value of accrued capital gains on stocks. Furthermore, gains attributed to common stock which was sold will be offset to a considerable extent by the gains which accrued to those individuals who reinvested the proceeds of their sales. Although households did have net equity sales of $2.7 billion in 1963, this would be expected to affect the true capital gains coefficient only slightly.[9] Thus for those individuals who provided data on their individual stock positions, the estimates of accrued capital gains are likely to be quite good.

In some cases no identification of specific shareholdings was given raising more serious questions of data reliability.[10] To guard against significant errors, for much of our analysis the sample used on owned homes; Y_N, normal income defined as the average of annual disposable income in 1962 and 1963;[12] $Y_T = Y_{63} - Y_N$; Z, 1963 accrued capital gains on stocks;[13] A_{-1} is net worth valued at market prices held at the end of 1962; and Age and Fam Size are continuous variables while Occu is represented by a dummy variable taking on the value 1 for households with self-employed heads and 0 for all other households.[14] In some regressions, V_{-1}, the net worth of stocks, and O_{-1}, the net worth of other assets, are substituted for A_{-1}.

The FRB survey did collect data on realized capital gains on all securities combined as a global figure for each household, but these figures seemed extremely weak and could not be checked against other data so that in the results presented in this paper no distinction was made between realized and unrealized capital gains.[15] No variable was introduced for accrued capital gains on assets other than stocks since such data were not available in the FRB survey. As a result, we were not able to check the intuitively implausible MPS result that the effect of an unexpected as well as expected change in value of homes, bonds and other non-equity assets on household consumption takes place immediately (i.e., in the same quarter). We were, however, able to check the implications of the MPS model for the timing pattern of consumption effects flowing from changes in the market value of stock owned.

Though theory does not permit us to prescribe the precise functional relation between S and Y_N, Y_T, Z and A_{-1} and the value of the associated parameters, we would expect that the capital gains coefficient (d) should be between 0 and the negative asset coefficient (e). For unanticipated capital gains, the coefficient would be expected to be close to zero; for anticipated gains, close to (e).[16] However, even though investors are likely to have some long-run capital gains rate in mind when they invest, short-run capital gains may be almost entirely unanticipated. Since any anticipated capital gains are presumably highly correlated with initial assets (and theoretically would be expected to have the same coefficient), the (d) coefficient in the saving regressions should reflect mainly the effect of unanticipated gains. We are interested, of course, in both the (d) and (e) coefficients, measuring respectively the short-run (same year's unanticipated capital gain) and long-run asset effects on saving.

Unfortunately, there is reason to believe that only the capital gains coefficient is measured with any precision, at least for the great bulk of households who specified the names and numbers of shares of individual stocks owned. The assets or net worth coefficient is likely to be substantially downward biased (in absolute value) by tastes effects, since households which have a high propensity to save will have both high savings and high assets. However, the capital gains coefficient is likely to be much less biased by tastes effects since tastes are much more highly correlated with assets than with capital gains. Indeed, when assets as well as capital gains are included in the savings regression, the correlation of capital gains and tastes should be close to zero.

Table 1 presents the major statistical results obtained from the analysis of the FRB survey, with data for each household weighted by the appropriate sampling ratio. The definitions of the saving, income, capital gains and asset variables used in these regressions are those specified by the FRB[17] except that depreciation on owned homes (estimated at

1.5 percent of beginning of year market value by the 1950 Survey of Consumer Finances) was deducted both from saving and assets. While this depreciation adjustment was made to improve conceptually the saving and wealth estimates, it should be noted that the capital gains coefficient was not affected appreciably by this adjustment. The most important remaining deficiency in the measures of saving and assets most commonly used in Table 1 is that the change in and the level of equity in insurance are generally not reflected in saving or assets. However, several regressions were computed in which the change in and the level of the surrender value of life insurance were added to the saving and asset variables, and one of these regressions is presented in the table. Equation (1) in the table, which covers all 2,159 households in the sample, points to a low depressant effect of capital gains (Z) and assets (A_{-1}) on saving, a rather high normal income (Y_N) effect, and a low transitory income (Y_T) effect. Only the Y_N and A_{-1} coefficients are statistically significant by the usual standards, and the goodness of fit ($\bar{R}^2$) for the equation as a whole is quite low even for un-grouped cross-section data.

Since there are, except for the respondents who give the names and quantities of individual shares owned, substantial errors in reporting the value of stock assets and in estimating capital gains on stock,[18] the two variables in which we are particularly interested, a second set of regressions, Equations (2) - (3), was based on those households which gave satisfactory detail on stock ownership. This second sample consisting of 1,770 households excluded those respondents reporting stock ownership who did not stipulate the amounts of individual stocks held accounting for at least 90 percent of the value of all stock reported owned by the household. The exclusion was designed to ensure a reasonable correspondence between reported and actual ownership of stock for those households reporting stock ownership. Equation (2) is directly comparable to (1), except for the smaller sample, and (3) substitutes the market value of stocks (V_{-1}) and other assets (O_{-1}) for total initial assets.

The results of Equations (2) - (3) represent an improvement over Equation (1), with a moderately higher overall correlation, and with more generally satisfactory regression coefficients. Thus the coefficients of Y_T are now statistically significant (at the customary 95 percent level) and, as would be expected, are higher than those of Y_N. Of special importance for our analysis, the coefficients of Z not only have the theoretically anticipated sign but are significant as well. Of the other economically important variables, the coefficient of A_{-1} is increased somewhat (in absolute value). The substitution of V_{-1} and O_{-1} for A_{-1} gives unreasonable results for the coefficient of O_{-1}, perhaps partly as a reflection of the multicollinearity between V_{-1} and O_{-1} and partly as a reflection of very large random errors in the measurement of O_{-1}. (Of the other socio-economic-demographic variables, none is significant though there is some indication that small families and entrepreneurs save more than other households.)

To eliminate other apparent sources of error in measuring saving, income, and assets, households which did not pass a number of quality checks were excluded resulting in a reduced sample of 1,195 households, and a new set of regression results is presented in Equations (4) - (6). Households were excluded if (1) net worth was negative, (2) normal income was less than $1,000, (3) the absolute level of saving was greater than normal income, and (4) stock assets at the beginning of year had a market value of less than \$200,000 and capital gains were in excess of the value of stocks held, or stock assets had a market value of \$200,000 or more and capital gains were in excess of half the value of stock held. The first of these exclusions was by far the most important and the fourth quite unimportant in reducing the sample size. The third is the most important exclusion in terms of impact on the regression coefficients. Without the first adjustment, the A_{-1} coefficients tend to be overstated. With out the second and third exclusions the Y_N coefficient tends to be overstated, while the third exclusion also serves to offset an overstatement in the Z and A_{-1} coefficients. However, empirically only the third exclusion has an appreciable effect on the coefficients.

Equations (4) - (5) for the 1,195 sample are the counterparts of (2) - (3) for the 1,770 sample. It will be noticed that the normal income and asset coefficients are substantially reduced, and the asset coefficient is no longer significant,[19] but that the capital gains coefficients are much less affected. (The occupational dummy variable is now significant or close to significant, with entrepreneurs again saving more than other households.) Equation (6) is the same as (4) except that savings and assets are now expanded conceptually to include the change in and level of the cash surrender value of private life insurance.[20] This provides an approximate measure of household saving which conforms more closely to the measure used in the aggregate national income statistics and econometric models. The adjustment of saving and assets for private life insurance does not substantially affect the regression coefficients either here or in other equations tested.

To eliminate one last potentially significant source of error in measuring assets and capital gains, households which reported owning no stock were excluded from the preceding sample of 1,195 households, resulting in a much smaller sample of 303 households. The reason for this exclusion is that validation checks suggest that a number of such households do in fact own stock.[21] If the families with stock but reporting none are relatively more numerous at the upper (lower) than at the lower (upper) end of the wealth and income scales, their inclusion tends to bias upward (downward) the estimates of the A_{-1}, V_{-1} and Z coefficients, and to bias downward (upward) the Y_N coefficient. On the other hand, while the quality of the sample is further improved, its size and representativeness are greatly reduced and the resulting sample is biased in favor of households which have a propensity for saving in the form of stock. Moreover, while the mean income, age of head, net worth and capital gains of households are not affected appreciably as the sample is reduced in size to 1,195 households, the mean age of head as well as net worth and capital gains (though not income) are increased substantially as the sample is further reduced to 303 households.

As a consequence, we are inclined to believe that the 1,195 sample results may be preferable to those obtained from the 303 sample, especially when the 303 sample households are weighted as in Equation (7) using the 128 weight classes which result from the FRB survey weighting procedures. In view of the small size of the remaining sample relative to the number of weight classes, obviously few or no observations are included in many of the weight classes if the FRB weighting procedures are followed. To handle this problem, the number of weight classes was reduced to twelve based upon normal income. The Y_N weights are internally generated by adding up the FRB weights of all the individual households within the larger normal income classes. Class sizes were chosen so as to contain an approximately equal number of observations.

Fortunately, when the Y_N weights were used in Equation (8), the statistical results were generally quite close to those obtained in the corresponding Equation (4) based on the 1,195 sample. The most noteworthy difference between Equations (7) and (8) or (4) is in the correlation, with $\bar{R}^2$ for (7) much higher than for the others. However, little confidence can be placed in (7) in view not only of the biased age composition of the sample but also the large weights ascribed to a small number of statistically biased observations in many of the 128 weight classes.

Finally, since the first eight equations in Table 1 all display significant heteroscedasticity in the residuals so that the t-values are overstated, four other equations derived from the 1,195 sample have been added (Eq. (9) - (12)) which correct for this problem by dividing S, Y_N, Y_T, Z, A_{-1}, and the constant term either by $A_{-1}^{.25}$ (Eq. (9) and (10)) or by $Y_N^{.25}$ (Eq. (11) and (12)). (This is equivalent to assuming that for different values of these independent variables the ratio of the original variance of residuals to $A_{-1}^{.5}$ or $Y_N^{.5}$ is constant.) Equations (9) and (11) like the first eight equations represent weighted regressions, while (10) and (12) are unweighted. The first three of these four equations almost completely eliminate heteroscedasticity in the residuals while the last, Equation (12), substantially reduces heteroscedasticity but it is still statistically significant. The estimated capital gains effect on saving in these four regressions is somewhat lower on average than in the corresponding Equation (4) which is weighted but not adjusted for heteroscedasticity, but the differences are probably not statistically significant.

Since age is not significant in the regressions in Table 1 in spite of the important role in saving attributed to it in life-cycle theories, Equation (4) which is based on the 1,195 sample was recomputed for each of five age groups and the results presented in Table 2. The capital gains coefficients show substantial variability, are never significant, and hence are not as satisfactory as the more aggregate results. However, all the capital gains coefficients have the correct sign and the variability may well reflect the small size of the sub-samples. When the sample households are classified into three instead of five age groups, the coefficient of capital gains for the middle age group, viz. 40-59 years for the age of household head, is statistically significant (with a coefficient of -.039 and t-value of 4.24).[22]

The estimation of other mathematical forms among the variables which appear in Tables 1 and 2, the substitution of other measures of Y_N (and hence Y_T), the substitution of normal and transitory labor income for total Y_N and Y_T, and adjustment for measurement error did not change the range of Z coefficients from those indicated above. The mathematical forms included dividing all variables (and the constant term) by Y_N, and, separately, adding square terms in some of the independent variables. An alternative measure of Y_N was obtained by regressing 1963 household incomes on the average incomes of a number of socio-economic-demographic groups to which the household belonged and then deriving the normal income of each household from this regression. As suggested earlier, measurement error would be expected to introduce some downward bias in the coefficient of Z, but this bias is likely to be small at least in the 1,195 and 303 samples.[23]

The question that remains to be answered is how the survey results on asset effects and especially on stock asset effects compare with the MPS and other time-series results. There are three relevant comparisons that can be made at least in theory; these relate to the coefficients of Z, V_{-1} and O_{-1}. However, as noted earlier, the regression coefficients of V_{-1} and O_{-1} are likely to be substantially biased downward (in absolute value) towards zero as a result of the impact of taste effects on the results of a single cross-section analysis. For this reason, the fact that the coefficients of V_{-1} and O_{-1} in the MPS model are substantially larger than those provided by the analysis of survey results is not significant evidence that the MPS results are biased upward. Such evidence could only be provided by panel cross-section data which can hold tastes constant.

The coefficients of Z in the analysis of 1963 survey results are likely to be little affected by tastes effects unlike the asset coefficients. The Z coefficients estimated from the survey regressions, ranging from -.015 to -.041, may be compared with corresponding estimates of -.040 for the last published version of the MPS model, -.032 from the latest unpublished version, and much lower figures from other time-series regressions.[24] Thus, the FRB survey results suggest that the latest MPS estimate of the first year's effect of a change in the value of stock assets on saving and consumption is not far off from the true figure though it may be a little on the high side, while the corresponding effects implied by most other time-series results seem too low. However, it should be noted that the time-series estimate of the first year's effect of a change in the value of stock assets on saving and consumption represents an average over a period of years while the survey results presented here are for a single year only.

To obtain a more definitive result for the first year's effect in general, survey data for other periods would be needed, and to obtain a complete answer to the long-term as well as short-term time-sequence of asset effects would require panel survey data. The statistical findings in this paper not only strongly support the inclusion of capital gains, and implicitly net worth, in the consumption function, but also indicate that survey data constitute an unusually promising source of information for quantifying the time-sequence of asset effects on saving and consumption.

Table 1

Household Savings Regressions for Different Samples

Reg. No.	Size of Sample	Constant	Y_N	Y_T	Z	V_{-1}	O_{-1}	A_{-1}	Age	Family Size	Occu.	$\bar{R}^2$
(1)	2,159	-0.21	0.24	0.14	-0.018			-0.014	-0.002	-0.119	1.08	.035
		(-0.27)	(5.84)	(1.08)	(-1.40)			(-8.23)	(-0.20)	(-1.30)	(1.53)	
(2)	1,770	-0.50	0.23	0.41	-0.040			-0.016	0.002	-0.100	0.85	.061
		(-0.59)	(5.01)	(2.73)	(-3.77)			(-9.80)	(0.18)	(-1.02)	(1.09)	
(3)		-0.21	0.19	0.47	-0.040	-0.021	0.000		-0.004	-0.085	0.33	.075
		(-0.25)	(4.03)	(3.16)	(-3.76)	(-11.18)	(0.05)		(-0.29)	(-0.87)	(0.42)	
(4)	1,195	-0.30	0.10	0.32	-0.028			0.001	0.001	-0.016	0.47	.099
		(-1.01)	(6.06)	(6.60)	(-4.85)			(1.20)	(0.25)	(-0.43)	(1.81)	
(5)		-0.38	0.11	0.30	-0.041	0.005	-0.005		0.003	-0.021	0.62	.112
		(-1.27)	(6.70)	(6.29)	(-6.31)	(3.84)	(-2.83)		(0.71)	(-0.54)	(2.38)	
(6)		-0.45	0.09	0.36	-0.032			0.002	0.004	-0.26	0.24	.084
		(-1.37)	(4.86)	(6.72)	(-4.98)			(2.01)	(0.81)	(-0.61)	(0.83)	
(7)	303	0.73	0.04	0.33	-0.015			0.001	-0.010	-0.004	1.42	.422
		(0.77)	(1.16)	(12.10)	(-1.30)			(0.48)	(-0.79)	(-0.03)	(2.39)	
(8)		-0.82	0.09	0.35	-0.021			0.001	0.009	0.067	1.44	.106
		(-0.97)	(2.64)	(3.93)	(-2.19)			(0.58)	(0.81)	(0.60)	(2.31)	
(9)	1,195	-0.40	0.11	0.24	-0.019			-0.002	0.002	-0.008	0.30	.094
		(-3.68)	(7.86)	(5.57)	(-1.05)			(-0.64)	(1.25)	(-0.56)	(2.69)	
(10)		-0.54	0.18	0.32	-0.029			0.000	0.001	-0.055	0.16	.134
		(-2.93)	(10.27)	(4.66)	(-4.88)			(0.32)	(0.54)	(-2.05)	(1.02)	
(11)		-0.41	0.11	0.29	-0.019			-0.001	0.002	-0.011	0.31	.087
		(-2.28)	(6.28)	(6.42)	(-2.10)			(-0.72)	(0.83)	(-0.55)	(2.32)	
(12)		-0.10	0.11	0.51	-0.029			0.002	-0.001	-0.026	0.14	.141
		(-0.21)	(4.80)	(7.19)	(-8.70)			(3.26)	(-0.19)	(-0.58)	(0.59)	

Note: S is saving for 1963 net of depreciation on owned homes and in Equation (6) is adjusted for saving in the form of life insurance; Y_N, normal income defined as the average of annual disposable income in 1962 and 1963; $Y_T=Y_{63}-Y_N$; Z, 1963 accrued capital gains on stocks; V_{-1}, the market value of corporate equities at the end of 1962; O_{-1}, the market value of assets other than equities at the end of 1962; A_{-1}, market value of assets or net worth held at the end of 1962; and Age and Family Size are continuous variables while Occu is represented by a dummy variable taking on the value 1 for households with self-employed heads and 0 for all other households. Figures in parentheses are t-statistics. Dollar variables are measured in thousands of dollars. $\bar{R}^2$ is R^2 adjusted for degrees of freedom.

Table 2

Household Savings Regression by Age of Head

Age of Household Head	Constant	Y_N	Y_T	Z	A	Age	Family Size	Occu.	$\bar{R}^2$
-34	-2.10 (-2.70)	.20 (4.57)	.38 (3.65)	-.165 (-.26)	.018 (1.81)	.048 (1.57)	-.034 (-.50)	.04 (.08)	.236
35-44	3.78 (1.68)	.04 (1.07)	.23 (2.18)	-.034 (-1.62)	-.001 (-.16)	-.092 (-1.76)	-.035 (-.39)	1.59 (3.00)	.098
45-54	.63 (.22)	.12 (3.12)	.33 (2.49)	-.039 (-.79)	-.003 (-1.19)	-.009 (-.15)	-.143 (-1.50)	.71 (1.28)	.049
55-64	-.18 (-.05)	.11 (3.02)	.20 (2.29)	-.003 (-.18)	.004 (3.36)	-.001 (-.02)	-.014 (-.13)	-.90 (-1.44)	.173
65+	.95 (.53)	.11 (2.47)	1.15 (6.00)	-.010 (-.89)	-.003 (-1.09)	-.013 (-.56)	-.065 (-.56)	-.84 (-.99)	.161

See Table 1 for explanation of symbols.

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National Quotation Bureau, OTC Industrial Stock Average, 1962-3.

New York Stock Exchange, Composite Stock Index, 1962-3.

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Survey of Current Business, Table 57, July 1966, 29.

Footnotes

[*]The authors are, respectively, Richard K. Mellon Professor of Finance and Economics and Ph.D. student in Economics, at the University of Pennsylvania. We wish to thank the National Science Foundation and the American Bankers Association for financial support and the Board of Governors of the Federal Reserve System for making previously unpublished survey data available to us.

[1]A description of these surveys and of the data collected appears in Projector and Weiss, and Projector.

[2]The latest published version of the MPS model appears in Modigliani, but a more recent version is available in Equations in the MIT-PENN-SSRC Econometric Model of the United States.

An evaluation of the forecasting ability of the MPS model compared with the well-known Wharton model, which indicates a relatively favorable record for the MPS model, is contained in Shinjo. An evaluation in Nelson of the forecasting ability of the MPS model compared with the best auto-regressive models mechanically derived from historical data is less favorable to the MPS model.

[3]E.g., see Crockett and Friend (1967).

[4]This updated version represents revisions incorporated as late as January 1973. As compared with the estimates in Modigliani, which did not break down wealth into stocks and other assets, the long-run asset coefficient (assumed to be the same for both forms of assets) did not change appreciably from the earlier estimate. However, a very much lower proportion of the effect on consumption of changes in the value of stock and a very much higher proportion of the impact of changes in the value of other assets take place in the first three months.

[5]See Friend and Shor.

[6]Ibid.

[7]A fairly close correspondence between reported and actual ownership for respondents who give the number of shares owned is indicated in validation tests discussed in Ferber, Forsythe, Guthrie and Maynes. In contrast a recent revision of the estimated market value of outstanding stocks in the U.S. made as part of the SEC Institutional Investor Study showed very little difference between the earlier SEC figures and the revised figures for the early 1960's but a $283 billion or 37 percent upward revision by 1968, most of the discrepancy occurring after 1964.

[8]Saving in the form of housing was adjusted for depreciation by deducting 1.5 percent of the initial value of housing as suggested by the 1950 FRB Survey of Consumer Finances. In that part of our analy-

sis where the survey definition of saving was expanded to include life insurance, the change in cash surrender value was used as a measure of saving in life insurance.

[9]The reduction in household equity holdings is from the Survey of Current Business (July 1966), p. 29. Total equity holdings at the beginning of 1963 are $527 billion as estimated by the SEC Institutional Investor Study. Accrued capital gains are estimated at $100 billion (see footnote 11).

[10]Ferber et al. found that in these cases the wealthy under-reported their holdings whereas the (relatively) poor overstated their holdings.

[11]The composite 1963 NYSE capital gain was 18.1 percent and the OTC gain was 20.4 percent as determined respectively from the New York Stock Exchange Common Stock and National Quotation Board indices.

[12]Normal income was also defined as the expected value of disposable income derived by regressing 1963 household income on the average incomes of each of a number of socio-economic-demographic groups to which the household belonged, with no important changes in the short-run asset effects presented here.

[13]Accrued gains in 1963 are gains reflecting the price changes during that year.

[14]Dummy variables were also introduced for age, distinguishing among the less than 39 years old, 40 to 59 years, and 60 and over

years head of household groups, without any noteworthy change in regression results.

[15]External sources as well as the survey data indicate that realized capital gains on stock are a small fraction of unrealized gains.

[16]For a theoretical discussion of the signs and values of these parameters, see Friend and Taubman and Crockett and Friend (forthcoming).

[17]See Projector.

[18]See Ferber *et al*. Even for respondents giving the number of shares of all stock reported owned, there is some tendency to overstate the size of very small holdings and understate very large holdings, and thus to bias somewhat upward (in absolute value) the coefficient of Z.

[19]In one equation, the stock assets coefficient (V_{-1}) is significant but has the wrong sign.

[20]It would be preferable to use equity instead of cash surrender value but the data are not available in the FRB survey.

[21]See Ferber *et al*.

[22]Additional tests examining the possible impact of age, the omission of age, expected versus unexpected capital gains, sensitivity of the results to sample size, and breakdown of the sample by normal income will be included in Lieberman. The omission of the age variable from either the five or three age group regressions has virtually no effect on the Z coefficients.

[23]The sensitivity of the Z coefficient to measurement error was tested by adding to Z a random component normally distributed with mean zero and standard deviation $k\sigma_z$ where k varied between .1 and .5 and σ_z, the standard deviation of capital gains, is $13,701. Five simulations of Eq. (4) of Table 1 were run for each value of k. For k equal to .1 the mean value of the Z coefficient was -.028 with a standard deviation of .00049. The corresponding figures were -.022 and .00148 for k = .4, and -.020 and .00163 for k = .5. At least for the 1,195 and 303 sample sizes, we believe the data to be measured with sufficient accuracy that k is likely to be well under .5.

[24]The estimate for the impact of unanticipated capital gains on saving in the MPS model was obtained by assuming that the MPS quarterly capital gains coefficients are equal to a weighted combination of the fully anticipated MPS wealth effect (.0544 in the latest version of the model) and the unknown unanticipated capital gains effects. The effect in subsequent periods of unanticipated capital gains in any quarter is the product of the weight of unanticipated gains in that quarter multiplied by the appropriate quarterly coefficient of such gains in the sequence of coefficients reflecting adjustment of households to the new level of assets. This formulation leads to five unknowns if the weight of unanticipated or anticipated capital gains is assumed constant (the four quarterly unanticipated capital gains coefficients and the weight of unanticipated--or anticipated--capital gains) but only four equations (one for each MPS quarterly coefficient). If the effect of contemporaneous capital gains is zero, an upper bound of .3 is found for the weight of anticipated capital gains in the latest

version of the MPS model. Assuming values between 0 and .3 for the weight of anticipated capital gains results in a narrow bracketed solution for the quarterly impact of unanticipated capital gains permitting use of the midpoints. The midpoints of the bracket for each quarter may then be applied to the percentage of 1963 capital gains which accrued during each quarter to estimate the impact of unanticipated capital gains on 1963 savings. A simple average of the NYSE and OTC composite index was used to calculate the percentage of 1963 capital gains which accrued during each quarter.