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# SOCIAL SECURITY AND CONSUMER SPENDING IN AN INTERNATIONAL CROSS SECTION*

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This paper expands the base of empirical evidence on the social security aggregate private saving issue by examining the behavior of consumer expenditure in 16 industrialized countries over the 1951–60 period. The results are mixed in that the time series movements of social security exhibit a positive relation to consumer spending, while the cross-sectional variations reveal a negative association. Our overall conclusion is that the cross-country evidence provides neither empirical support for the hypothesis that social security depresses private saving nor an empirical refutation of that hypothesis. We argue that this indeterminacy of results applies also to previous studies of U.S. time series and to analyses of household cross sections in the U.S.

## 1. Introduction and theoretical considerations

The theoretical argument for a downward effect of a 'pay-as-you-go' social security program on private saving was presented by Feldstein (1974) in the context of a 'life-cycle' model. Individuals view anticipated social security benefits during retirement as a substitute for own preretirement savings and are therefore motivated to diminish their accumulation of assets during working years. However, this conclusion emerges only because the model assigns the government a monopoly position in respect to intergenerational transfers. In fact, most individuals have numerous private opportunities for shifting income across generations. Parents make voluntary contributions to children in the form of educational investments, other expenses in the home including parental time, and bequests. Children – especially before the expansion of social security – provide support for aged parents. To the extent that private, voluntary transfers of this sort are operative – and casual observation suggests that such transfers in the appropriate broadly defined

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sense are pervasive – the main response to more social security, i.e. to more governmentally imposed intergenerational transfers, would be a shifting of private transfers by an amount sufficient to restore the balance of income across generations that was chosen previously. An important effect of this type is the apparently strong influence of social security in reducing the fraction of retired people who live with, and presumably receive support from, their children.[1] When this sort of private offset to social security occurs, the downward effect on private saving would no longer be predicted.

A full theoretical analysis of social security would consider a variety of other issues, such as the program's insurance aspects (in comparison with private alternatives that would include the role of the extended family), the significance of childless individuals who are likely to show small concern for particular members of later generations (but, who must be offset by other individuals with an above-average number of children), the responsiveness of retirement and work intensity to eligibility rules for receiving benefits, and the implications of social security as an element of the overall distorting influence of non-lump-sum governmental taxes and transfers. Although these considerations are of some importance, they have ambiguous implications for private saving and capital accumulation and, in any event, do not yield the dramatic, first-order saving effects that emerge from the simple life-cycle model.

Overall, economic theory provides neither an a priori argument for a strong depressing effect of social security on saving nor decisively rules out an important influence. The crucial issues are empirical.

In a study of aggregate time series data for the U.S. since 1929, Feldstein (1974) concluded that 'social security depresses personal saving by 30–50%'. A re-examination of the evidence [Barro (1978a)] indicated that this conclusion was unwarranted, although the results were not sufficiently strong to rule out decisively an economically important downward effect of social security on private saving. A survey of the U.S. time series evidence is contained in Esposito (1978) and a further discussion of this material will appear in a 1979 issue of the *Social Security Bulletin*.

Especially because of the indeterminacy of the U.S. time series studies, it is important to consider other forms of evidence. The present paper expands that base of evidence by examining the behavior of consumer expenditure in a time series/cross section sample of 16 industrialized Western countries over the 1951–60 period. [A related body of evidence has been considered cross-sectionally in terms of time-averaged data by Houthakker (1965); Modigliani (1970); Feldstein (1977); and Sterling (1977).] The last two studies also include a social security variable. The main conclusion from our in-

[1]See Munnell (1974, ch. 2). For a discussion of the private offset to social security see Barro (1974, 1978b) and Drazen (1978).

vestigation is that – as in the U.S. time series case – the evidence does not support an inverse effect of social security on private saving. However, the problems with data and estimation are more pronounced in the present case than in the U.S. time series analysis. Hence, the conclusions are, unfortunately, again of an indecisive nature. The cross-country evidence examined provides neither a basis for believing that social security reduces private saving nor a basis for firmly discarding that belief.

## 2. Setup of the cross-country empirical framework

The basic idea of using a cross-country data sample in the present context is to exploit a greater variety of experience with social security programs than is afforded by the time series evidence within a single country. In addition, cross-sectional observations avoid some problems involved with disentangling the effects of social security from those of other variables in a time series context. However, some serious identification problems arise in the cross-section context because of a possible two-way causation between private saving and social security expenditure [as noted in Aaron (1967, p. 19)]. The difficulties with constructing a homogeneous measure of social security and, to a lesser extent, of other variables are also more pronounced across countries than for a time series within a single country. This data problem has limited the present inquiry to the experience of 16 industrialized Western countries. (See table 1 for a listing of countries – a detailed data appendix is available on request.) Even for these 16 cases the social security measures are open to serious question with respect to conceptual consistency across countries. Furthermore, the sample is not ideal in terms of the range of cross-country experience since less developed (and even some industrialized countries) could not be included. Surprisingly, the data compiled by the International Labor Office on supposedly homogeneous measures of social security for the 1951–60 period seem more suitable for our purposes than data we have located for more recent years. Therefore, our present study does not utilize any post-1960 data. However, we have fully utilized the available ten-year time series experience for each country,[2] rather than limiting attention to a single time-averaged observation from each case.

The form of the consumer spending equation that is employed in this study is a simple linear relation that is based on earlier theoretical models. [See Modigliani and Brumberg (1954); Ando and Modigliani (1963); Modigliani (1966, 1970); Feldstein (1974, 1977); and Barro (1978a).] The functional form is

$$(C/Y)_{it}=(\alpha_0)_i+\alpha_1(G/Y)_{it}+\alpha_2\rho_i+\alpha_4(Y_{t-1}/Y_t)_i+\alpha_5(OLD)_{it}+\alpha_6(SS)_{it}+\varepsilon_{it}, \tag{1}$$

[2]Because observations are missing for a few years for some countries the total number of observations turns out to be 152.

where annual observations run over countries $i=1,\ldots,16$ and observations apply in most cases for $t=1951,\ldots,1960$. The variables, tabulated and defined in detail in the data appendix that is available from the authors, are consumer expenditure, $C$ (with no attempt at adjustment for consumer durable purchases or investment in human capital); gross domestic product, $Y$; government purchases of goods and services, $G$; the long-term growth rate of real per capita GDP, $\rho$ (assumed to be invariant over time for a given country); an unemployment rate variable normalized in an attempt to compensate for intercountry differences in measurement practices (see the notes to table 1 and the data appendix), $U$; the ratio of population over 65 to the total population, $OLD$; current social security benefits paid on old age, survivors and disability programs relative to population over 65 and divided by per capita GDP, $SS$; and a random error term $\varepsilon$, discussed below.

As eq. (1) is written it implies a homogeneity condition that a doubling of all income and expenditure variables ($Y_t$, $Y_{t-1}$, $G$, and social security benefits) leads to a doubling of consumer expenditure, i.e. the consumer spending-income ratio is independent of income scale. The rationale for this specification is developed in Modigliani and Brumberg (1954, p. 396, ff.). However, as noted by Feldstein (1977, p. 16), if higher income is associated with a lengthening of the relative amount of time spent in retirement, an increase in real income per capita would raise the desired ratio of wealth to income and thereby (for a given positive growth rate of real income) reduce the steady-state value of the consumer spending-income ratio. This effect can be explored by entering the variable $1/Y$, where $Y$ represents real per capita GDP, into eq. (1), and checking for a positive effect on $C/Y$. However, unlike the other variables that appear as ratios in eq. (1), the $1/Y$ variable requires intercountry comparability of real per capita income levels. Empirically, this measurement relies on the application of official exchange rate figures and the U.S. price index to each country's reported values of nominal per capita GDP. Because of its reliance on exchange rates, the $1/Y$ variable is subject to an extra element of measurement error.

The income concept used in the present study is gross domestic product (GDP). Because of the measurement problems with depreciation allowances and net factor income from abroad, it seemed that this "income" measure would be more homogeneous across countries than some alternatives. The present definition also encompasses household and business income, as seems appropriate.

The effect of government purchases (included in GDP) on consumer expenditure is examined by inclusion of a separate government purchases of goods and services variable, $G/Y$. The general role of this variable, i.e. the nature of the $\alpha_1$ coefficient in eq. (1), is viewed along the lines discussed in Bailey (1971, ch. 9). For this purpose, we abstract here from distinctions between permanent and temporary values of government purchases or private sector income. Suppose that government purchases $G$ are perceived by the representative member of the private sector as equivalent to $a_1 G$ units of private consumption and $a_2 G$ units of private saving. Hence, the private sector would augment its disposable income by including the imputation, $(a_1 + a_2)G$, for the value of government purchases. The condition $(a_1 + a_2) \gtreqless 1$ reflects the relative efficiency of government activities, where the present illustrative analysis does not allow for diminishing marginal returns to government expenditure. Suppose that the marginal propensity to consume out of augmented private sector disposable income is $\mu$, i.e. in a linear form

$$C^* = \text{constant} + \mu Y^*,$$

where $C^* = C + \alpha_1 G$ is augmented private consumption and $Y = GDP - G + (a_1 + a_2)G = GDP - (1 - a_1 - a_2)G$ is augmented private sector disposable income.[3] Therefore, after some substitutions, measured private consumption expenditure depends on GDP in accordance with

$$C = \text{constant} + \mu \cdot GDP - [(1-\mu)a_1 + \mu(1-a_2)]G. \tag{2}$$

The $\alpha_1$-coefficient in eq. (1) then corresponds to $-[(1-\mu)a_1 + \mu(1-a_2)]$. If $0 < a_1 < 1$, $0 < \mu < 1$, and $a_2 > 0$ then $-1 < \alpha_1 < 0$. The magnitude of $\alpha_1$ is smaller the larger the saving content of government purchases, $a_2$, the smaller the consumption content, $a_1$, and, if $a_1 + a_2 < 1$ so that more $G$ implies less augmented private sector disposable income, the smaller the marginal propensity to consume, $\mu$.

The effect of the long-term growth rate of real per capita income, $\rho$, is discussed theoretically in Modigliani (1966, p. 167). Because the anticipation of growing income per head implies that households who are currently positive savers (the younger households) will wish to provide for more consumption during retirement than the amount enjoyed by currently negative savers (the older households), there tends (though not inevitably) to be a negative association between $\rho$ and the consumption-income ratio, so that $\alpha_2 < 0$ in eq. (1). Population growth, which should have a separate negative effect on $C/Y$ [Modigliani (1966, p. 166)], turns out to be unimportant empirically, as was also true in the results reported by Modigliani (1970, p. 211).

The unemployment rate (used previously in time series studies of saving behavior by Ando and Modigliani (1963, p. 61); Feldstein (1974); and Barro (1978a)] and the lagged income variable are cyclical measures intended to account for any differences between current and permanent income. Since, for a given value of $Y_t$, $U_t$ and $Y_{t-1}$ would both tend to be positive predictors

[3] This analysis treats the government deficit as equivalent to current taxation in terms of consumption effects.

of future income, the $\alpha_3$ and $\alpha_4$ coefficients should both be positive. It turns out in the present empirical analysis that the lagged value of income is unimportant once the unemployment rate is included in the equations.

The old age variable (population over 65 relative to total population), which is similar to variables used by Modigliani (1970, p. 213 – the $R/W$ variable) and Feldstein (1977, p. 14 – the $AGE$ variable), is more difficult to interpret. If the variable proxies for the fraction of the population that is retired (and, hence, in a steady state, also for the fraction of planned time spent in retirement), then there would be a long-run positive relation of the $OLD$ variable to the desired ratio of wealth to income. Therefore, for a given positive growth rate of real income, the effect of $OLD$ on $C/Y$ ($\alpha_5$) would be negative. However, for a given set of saving plans and a given steady-state age distribution, an increase in the current ratio of retired persons to total population would increase the fraction of people who have a high propensity to consume, so that $\alpha_5$ would be positive. An additional problem of interpretation is that the old age variable need not proxy for the fraction of retired persons, since increased life expectancy and the associated improvement in health throughout the life cycle would tend to raise both the working and retirement spans. In a sense the theoretical indeterminacy surrounding the $OLD$ variable may be an advantage since it makes it easier to rationalize the confusing results on this variable that appear in the subsequent empirical analysis.

The social security variable is included to test the basic hypothesis that a greater expected retirement benefit from this source would increase $C/Y$. The social security measure (see the data appendix for details) is based on current social security benefits relative to the total population over 65, divided by per capita GDP. Data on amount of benefits (under compulsory old age, survivors and disability programs, although the specific concepts vary across countries) were obtained from issues of the International Labor Office's *The Cost of Social Security* and the *Yearbook of Labor Statistics*. The form of the variable assumes that real benefits per recipient and the fraction of the old age population covered by social security enter multiplicatively in affecting real per capita consumer expenditure.[4] Feldstein (1977, pp. 23 ff.) measures these two aspects of the social security programs separately, although the data do not allow this separation in an entirely satisfactory manner. In any event his results (tables 1 and 4) are consistent with the multiplicative form of these variables.

Feldstein (1977, table 1) adds as a separate explanatory variable in parts of

[4]Real per capita consumer spending is assumed to depend on real social security benefits relative to the population over 65. The latter ratio equals average real social security benefits per covered person times the ratio of covered persons to the total population over 65. The $SS$ variable, as defined in the text, emerges when real per capita consumer spending is divided by real per capita GDP to obtain the $C/Y$ variable on the left side of eq. (1).

his study the labor force participation rate of the aged. Since this variable tends to hold constant the induced retirement effect of social security, the $SS$ variable would then reflect – as he points out – only the hypothesized positive 'wealth' effect on $C/Y$ (offset by the adjustment of private intergenerational transfers). This type of labor force participation rate variable has not been included in the present analysis for two reasons. First, our examination of the data indicates that measurements of labor force participation by age are available for most of the included countries during the 1950s for no more than a few years. Furthermore, the concepts of labor force participation seem to differ widely across countries. Secondly, our principal interest is in the net effect of social security on consumption, rather than in the effect with labor force participation held constant. (Although the effects of social security on retirement, hours of work, etc. are of substantial interest for their own sake.)

In estimating relations of the form of eq. (1) for a cross-country sample, Houthakker (1965), Modigliani (1970) and Feldstein (1977) use a weighting scheme that assumes that the variance of the error term, $\varepsilon_{it}$, is proportional to the reciprocal of population, $1/POP_{it}$. This variance property would hold if each country-wide observation represented an average of $POP_{it}$ independent units,[5] with a homogeneous variance at the unit level. On the other hand, if observations also include country-wide components in the error term, then the variance may decline less rapidly with population. Our empirical analysis (footnote 11 below) suggests that the error variance does decline with population, although not quite as rapidly as the (negative) unitary elasticity suggested above. However, estimations with a weighting scheme that assumes the unitary elasticity specification turn out to yield results that are close to those based on internally-estimated weights. We report results below in both weighted and unweighted forms, although the weighted estimates would seem to be more reliable.

Finally, eq. (1) has been estimated with a common intercept, $\alpha_0$, across countries and with individual country intercepts. (The latter estimation would, of course, be impossible if a single time-averaged observation had been used for each country.) The common intercept procedure would use both the cross-sectional and time series variations in the data to estimate the effects on $C/Y$ of social security and the other variables. On the other hand, when individual intercepts are admitted, the coefficient estimates would be based on an average across countries of time series variation, rather than on cross-sectional variation in mean country characteristics. For example, the effect of the long-term growth rate variable cannot be estimated at all in the

[5]More plausibly, the number of independent units would be some fraction of $POP_{it}$. If each unit of population were independent, then even a large error variance at the unit level would become a negligible variance in country-wide observations that represented averages over a few million units.

individual intercept case. Ideally, from the standpoint of using the common intercept form, the hypothesis of equal intercepts across countries would be in agreement with the data. Unfortunately, this equal-intercept hypothesis is decisively rejected in our sample with the current list of explanatory variables. The single-intercept form could still be viewed as applicable to a situation where the error term includes an element that varies across countries but is invariant over time within a single country.[6] In any event – as discussed below – we presently end up with two types of parameter estimates: first, estimates from common-intercept equations that provide precise-looking estimates of coefficients that resemble in a number of respects – though not for the social security variable – the estimates from previous cross-country saving studies. However, these estimates show a relatively poor fit and a tendency for individual countries to have either all positive or all negative residuals over time. Both indicators of statistical deficiency seem to reflect the underlying significant differences in individual country intercepts. Second, we obtain estimates with individual intercepts that show a good fit and only moderate positive serial correlation of residuals (for individual countries over time), but which also show a pattern of estimated coefficients that differs in some major respects from those obtained by the first procedure.

## 3. Empirical results

The basic empirical results are contained in table 1. The data base comprises 16 countries with 152 total observations. The first four columns of the table relate to weighted regressions in which observations on all variables are premultiplied by the corresponding value of the square root of population. The weighting scheme is discussed in footnote 11 below. The last four columns of the table deal with unweighted regressions. Columns 1, 3, 5 and 7 are in the homogeneous form shown in eq. (1), while the other four columns add the $1/Y$ variable. Columns 1, 2, 5 and 6 use a common intercept term, while the other columns use individual country (16) intercepts. (Note that the long-term growth rate $\rho$ is, by necessity, dropped from these last calculations.)

Consider, first, the results in the common-intercept form (columns 1, 2, 5 and 6). The coefficient of the government expenditure variable, $G/Y$, is negative as expected (recall that $Y$ is measured by gross domestic product), with coefficient estimates in the weighted form (columns 1 and 2) of $-0.6$ for the homogeneous specification and $-0.3$ for the nonhomogeneous form. These coefficients are somewhat lower in magnitude than expected, in the sense that they require a high saving content of government expenditure ($a_2$) in order to be consistent with the model described above in eq. (2). The old

[6]However, the estimation procedure should then be altered to take account of the cross-country structure of the error term.

age variable has a positive effect with a magnitude in the weighted form (columns 1 and 2) that is on the same order[7] as that reported by Modigliani (1970, p. 211 – dividing his coefficient by 100 and changing the sign, since his dependent variable is the saving ratio) and Feldstein (1977, table 4 – again changing the sign). The estimated coefficient on the long-term real per capita income growth rate, $\rho$, is negative with a magnitude in the weighted form (columns 1 and 2) that is similar to that reported by Modigliani (1970, p. 211 – the $y'$ variable with the opposite sign). Feldstein's (1977, table 4) results, which apply to the growth rate of total real income, are somewhat higher in magnitude [corresponding, however, to differences in results for total versus per capita real income growth rate variables reported by Modigliani (1970, p. 211 – the $y'$ and $y$ variables)].

With respect to the cyclical variables, which were excluded in the earlier cross-country saving studies that did not involve time series variation, the unemployment rate variable is positive as expected and significantly different from zero. This result accords with earlier findings for the U.S. time series in Barro (1978a).[8] The lagged income variable does not have an important effect.

The nonhomogeneity term, $1/Y$, is positive and significant (columns 2 and 6), indicating some negative effect of income level on the consumer spending–income ratio. Aside from an improvement in the fit, the main effects of the addition of this variable are a reduction in estimated intercepts (especially in the individual intercept cases below), an increase in the magnitude of the estimated growth rate coefficient, and a decline in the magnitude of the estimated $G/Y$ coefficient.

With respect to the variable of greatest current interest, $SS$, the coefficient estimates are significantly negative! i.e. more social security is estimated to imply less consumer expenditure. However, this result – which appears strongly in the common-intercept form of the consumer spending equations – does not hold up when individual country intercepts are introduced.

We have been unable to reconcile our results on social security from the common-intercept form of the equations with Feldstein's findings. His equations that are comparable to ours in the sense of excluding a labor force participation variable [Feldstein (1977, table 4)] report a negative effect of his social security variable ($BPA/Y$) on the *saving* rate. Since Feldstein used a single time-averaged observation for each country in an equation with (necessarily) a common intercept, we would have expected him to find a positive saving effect. The interpretation of Feldstein's results is also a puzzle

[7]For comparative purposes, the coefficient estimates in table 1 should be inflated by about 20% to adjust for differences in income concept – GDP versus private sector disposable income.

[8]As the $U$ variable is presently measured (see the notes to table 1), a unit change corresponds to a shift by one standard error in the unemployment rate. For the United States this change amounts to 1.25 percentage points in the actual rate. The corresponding change is substantially smaller for many of the other countries, because of the smaller historical volatility of their measured unemployment rates. See the data appendix for details.

Table 1

Regression results for consumer expenditure.[a]

| | Weighted by population | | | | Unweighted | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Constant | 0.76(0.06) | 0.62(0.05) | | | 0.58(0.07) | 0.54(0.07) | | |
| $G/Y$ | −0.61(0.08) | −0.29(0.09) | −0.04(0.09) | −0.09(0.08) | −0.43(0.14) | −0.29(0.14) | −0.14(0.16) | −0.15(0.15) |
| $OLD$ | 0.90(0.16) | 1.09(0.15) | −2.68(0.52) | −0.20(0.54) | 0.67(0.21) | 0.65(0.21) | −2.54(0.48) | −1.41(0.55) |
| $\rho$ | −0.71(0.21) | −1.19(0.20) | | | −0.30(0.29) | −0.70(0.32) | | |
| $U$ | 0.010(0.002) | 0.007(0.002) | 0.007(0.001) | 0.003(0.001) | 0.008(0.003) | 0.006(0.003) | 0.006(0.001) | 0.005(0.001) |
| $Y_{t-1}/Y_t$ | −0.06(0.05) | −0.01(0.04) | −0.04(0.02) | 0.01(0.02) | 0.09(0.07) | 0.10(0.06) | 0.04(0.03) | 0.03(0.02) |
| $SS$ | −0.12(0.02) | −0.09(0.02) | 0.05(0.02) | 0.07(0.02) | −0.08(0.03) | −0.08(0.03) | 0.01(0.03) | 0.05(0.03) |
| $1/Y$ | | 0.044(0.007) | | 0.095(0.013) | | 0.036(0.012) | | 0.064(0.017) |
| Individual intercepts: | | | | | | | | |
| Australia | | | 0.89(0.05) | 0.58(0.06) | | | 0.82(0.05) | 0.69(0.06) |
| Austria | | | 0.96(0.06) | 0.52(0.08) | | | 0.90(0.06) | 0.69(0.08) |
| Belgium | | | 1.05(0.06) | 0.65(0.08) | | | 0.98(0.06) | 0.80(0.07) |
| Canada | | | 0.86(0.05) | 0.59(0.05) | | | 0.80(0.04) | 0.69(0.05) |
| Denmark | | | 0.98(0.05) | 0.62(0.07) | | | 0.92(0.05) | 0.75(0.07) |
| Finland | | | 0.83(0.04) | 0.55(0.05) | | | 0.77(0.04) | 0.64(0.05) |
| France | | | 1.01(0.06) | 0.61(0.08) | | | 0.95(0.06) | 0.77(0.07) |
| Italy | | | 0.94(0.05) | 0.55(0.07) | | | 0.88(0.05) | 0.69(0.07) |

| | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Luxembourg | | | 0.88(0.07) | 0.52(0.07) | | | 0.82(0.05) | 0.66(0.07) |
| Netherlands | | | 0.85(0.05) | 0.50(0.06) | | | 0.80(0.04) | 0.63(0.06) |
| New Zealand | | | 0.92(0.05) | 0.59(0.06) | | | 0.86(0.05) | 0.71(0.06) |
| Norway | | | 0.89(0.06) | 0.53(0.07) | | | 0.83(0.05) | 0.66(0.07) |
| Sweden | | | 0.94(0.06) | 0.57(0.07) | | | 0.88(0.05) | 0.71(0.07) |
| Switzerland | | | 0.96(0.06) | 0.62(0.07) | | | 0.90(0.05) | 0.74(0.06) |
| U.S. | | | 0.90(0.05) | 0.62(0.06) | | | 0.84(0.05) | 0.72(0.06) |
| W. Germany | | | 0.88(0.05) | 0.51(0.07) | | | 0.84(0.05) | 0.65(0.07) |
| $R^2$ | 0.50 | 0.61 | 0.91 | 0.94 | 0.20 | 0.24 | 0.91 | 0.92 |
| D.W. | 0.4 | 0.3 | 1.4 | 1.3 | 0.2 | 0.2 | 1.5 | 1.4 |
| $\hat{\sigma}$ | 0.021 | 0.019 | 0.010 | 0.008 | 0.036 | 0.035 | 0.013 | 0.012 |

[a]*Notes:* The dependent variable in all regressions is the ratio of consumer expenditure to gross domestic product, $C/Y$. Constant is the common intercept term estimated for the equations in columns 1, 2, 5 and 6. Individual intercept terms are indicated for each country for the equations in columns 3, 4, 7 and 8. $G/Y$ is the ratio of government purchases of goods and services to gross domestic product. *OLD* is the ratio of population over 65 to total population. $\rho$ is the average real per capita growth rate of GDP calculated for each country over a period that ranged between 23 and 25 years. $U$ is the normalized unemployment rate variable: $(u-\bar{u})/S_u$, where $u$ is the measured rate, $\bar{u}$ is the mean of this rate and $S_u$ is its standard deviation for each country over the 1951–60 period. $Y_{t-1}/Y_t$ is the ratio of lagged to current real per capita GDP. *SS* is current real social security benefits paid relative to the population over 65, divided by real per capita GDP. $1/Y$ is the reciprocal of real per capita GDP, expressed by means of the official exchange rate and U.S. price index for each year in terms of constant (1970) U.S. dollars.

Detailed definitions and a tabulation of variables appear in the data appendix that is available from the authors.

Numbers in parentheses beside each coefficient estimate are standard errors. The calculations reported in columns 1–4 weight each squared residual in the least squares estimation by the corresponding value of population. $R^2$ for these cases takes account of the weighting scheme. D.W. is the Durbin–Watson statistic, which takes account here of the gap in sequence in movements from one country to the next. $\hat{\sigma}$ is the standard-error-of-estimate. For the weighted cases the $\hat{\sigma}$ value applies to a country with the mean value of population.

to Sterling (1977, p. 61), who used a single time-averaged observation for each country.[9] Sterling's conclusions (pp. 50, 62) are basically in accord with ours – in particular, with labor force participation behavior *not* held constant his estimates indicate a positive effect of social security on saving.

It is clear from table 1 that the fits of the common-intercept equations (columns 1, 2, 5 and 6) are not very good – in the case of the weighted equation that includes the $1/Y$ variable (column 2) the (weighted) $R^2$ is 0.61. The Durbin–Watson Statistics[10] are unsatisfactory, with values ranging between 0.2 and 0.4. It turns out that these statistics are more an indicator of persistent differences in levels of $C/Y$ across countries than of positive serial correlation of residuals over time within a single country.

Columns 3, 4, 7 and 8 of table 1 provide estimates with individual intercept terms. The hypothesis that these intercepts are equal across countries can be strongly rejected in all cases. For example, the statistic for the weighted case that includes the $1/Y$ variable (columns 2 and 4) is $F^{14}_{131} = 47$, 5% critical value = 1.7. As a related matter, the fit improves considerably when individual intercepts are admitted, with the $R^2$ now exceeding 0.9 and the Durbin–Watson statistics now lying between 1.3 and 1.5.[11]

Except for the cyclical and $1/Y$ variables it is difficult to interpret the coefficient estimates from the individual intercept cases. The government expenditure variable is now insignificant, while the old age variable changes sign and becomes sensitive to the inclusion or exclusion of the $1/Y$ variable. Most importantly, the coefficient of the social security variable also changes sign – now indicating a positive effect on $C/Y$ that differs significantly from zero except in the unweighted, homogeneous case (column 7). The estimates now seem comparable (after being increased by about 20% to account for differences in income measures, see footnote 7) to those reported by Feldstein (1977, table 4).[12] This comparability is actually surprising since Feldstein's

[9]Sterling reports results for a sample of industrialized nations and also for a much larger group of countries.

[10]These statistics are calculated taking account of the break in the dating of the sample in moving from one country to the next. Essentially, the data are used to estimate the autoregressive equation for the residuals, $\varepsilon_t = \rho\varepsilon_{t-1} + v_t$, where $v_t$ is a white-noise error term, with one observation dropped for each country in the sample. This procedure is valid under the null hypothesis that $v_t$ is identically distributed for each country.

[11]The relation between error variance and population can be examined through an iterative procedure starting with a regression of the squared residuals from each unweighted equation (columns 5–8) on population. The implied weights from this regression are used to re-estimate the $C/Y$ equation, and the resulting squared residuals allow for a second-round estimate of the weights. The result of these calculations is an estimated elasticity of error variance with respect to population of about $-0.7$ from equations based on the forms in columns 5, 6 and 8 of the table, and $-0.5$ for the column 7 equation. The 'standard errors' for these values are between 0.10 and 0.15, so that the estimates are 'significantly' lower in magnitude than the value of $-1.0$ used in the weighted equations in columns 1–4. However, the differences in estimates produced by these changes from the weighting structure used in columns 1–4 are of little consequence and it did not seem worth reporting them separately.

[12]It is also necessary to adjust for sign since Feldstein's dependent variable is the saving rate.

results are, as would be expected, otherwise comparable to the estimates from the common-intercept form of our equations and not comparable to the estimates from the individual intercept (time series-based) equations. However, for the common-intercept cases (columns 1, 2, 5 and 6 of table 1), our estimates show a significantly negative effect of social security on $C/Y$.

It seems that the flipping in sign of the social security coefficient (and perhaps also of the old age coefficient) with the change in treatment of the intercept term would relate to differences in the nature of cross-sectional and time series movements in the independent variables. The suggestion from the results is that the time series movements of social security (reflected in the individual intercept cases) involve a positive relation to $C/Y$, while the cross-sectional variations involve a negative association.

The explanation for the differences in these two types of relationships is unclear to us, although reverse causation from consumption propensities to the level of social security expenditure may be involved. To account for the results it is first necessary to have differences between the cross-country and time series responses of social security expenditure to autonomous shifts in consumption. The cross-country response would tend to be more important because these differences in consumption behavior seem to be more 'permanent' (as evidenced by the differences in intercept values across countries) than those occurring over time for a particular country. The second element of the reverse-causation explanation for our common-intercept results would have to be that countries that were inclined to consume a larger fraction of income would also be less likely to adopt generous social security programs. This association would arise if the forces that lead to high private saving lead also, through the political process, to generous public 'saving' programs. Aaron (1967, p. 19), in a study that focuses on cross-country determinants of social security spending, has also noted the possibility of two-directional causation between saving propensities and social security expenditure. However, he argues from the standpoint of adequate provision for life-cycle risks that low-saving countries would tend to spend more on social security. In any event, the principal conclusion from our present results seems to be that any desired sign for the social security variable in a cross-country consumer expenditure equation can be picked by judicious choice of specification, particularly with regard to common or individual intercept terms.

## 4. Conclusions

The overall conclusion from this study must be an uninspiring negative one – that the cross-country evidence does not provide empirical support for the hypothesis that social security depresses private saving and also does not permit an empirical refutation of that hypothesis. In this respect the results are analogous to those from the U.S. time series that were discussed above.

A final body of evidence that has been considered in previous research is a cross-section of individual households at a point in time in the U.S. [Munnell (1976); Feldstein and Pellechio (1977)]. The basic finding in these studies that is relevant in the present context is that an increase in prospective social security benefits reduces private asset accumulation during working years. While this result may be correct (there are some difficulties in isolating independent variation in the social security variables in the samples), it does not bear directly on the central issue, which concerns the impact of an overall social security program on aggregate saving and capital accumulation. The individual cross-section findings correspond to a positive effect of relative social security benefit (net of tax) positions on relative consumption – a relation that is consistent with the plausible hypothesis that more individual income means more individual consumption. The theoretical position for no aggregate saving effect of social security, which is based primarily on the view that voluntary private intergenerational transfers offset the government's actions, corresponds to the proposition that it is only one's relative social security benefits and tax position – and not the absolute level of social security – that produces shifts in consumption. A change in the scale of the program increases benefits and liabilities by equal amounts (if the benefits and liabilities of descendants are fully counted) and thereby has no effect on consumption. A cross-section of individuals at a point in time holds fixed the scale of the overall social security program, while examining only the effect of changes in individual relative positions. Therefore, these data provide no variation in the pertinent variable – the scale of the overall program – which is essential for tests of propositions that concern aggregate saving effects.

Our general assessment of present empirical knowledge is that, either in terms of individual components of evidence or in terms of the overall picture, there is no support for the proposition that social security depresses private saving. The effect of social security on saving and capital formation remains an open empirical issue.

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