ESTIMATING THE EFFECTS OF SOCIAL SECURITY REFORM: AN EULER-EQUATION APPROACH

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Abstract

This paper shows theoretically and empirically that an aggregate Euler equation relates the real interest rate to the growth rate of per capita consumption and the ratio of private and social security wealth plus asset income to consumption. Using the estimated Euler equation, the paper then calculates the steady-state effects of social security reform. Reforms that reduce the ratio of social security wealth to consumption are found to shift the balanced growth paths for the capital stock, output, and consumption upward appreciably.

JEL Codes: E21, E62, H31

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I. INTRODUCTION

The reform of the US social security system is now attracting a great deal of attention. A wide range of proposals for reform have been put forward, ranging from modest modifications to complete privatization.[1] In choosing among these reforms, it is important to quantify how they would affect the aggregate economy. A given reform is presumably more likely to command support, the larger are its favorable effects on net.

In recent years, economists have typically quantified the effects of social security reform in large computational general-equilibrium models based on the life-cycle hypothesis; e.g., Laurence Kotlikoff, Kent Smetters, and Jan Walliser (1998). Although the calibration of these models does make use of data, the models are not designed to fit the aggregate data in any very well-defined sense. Rather, the life-cycle structure of the model is simply imposed, and the parameters are chosen in order to match selected moments of the aggregate data. The life-cycle hypothesis is not the only possible model of household, however. One might therefore reasonably doubt the empirical validity of effects calculated in such a model.

An older literature attempted to estimate the effects of social security directly. The seminal study in this literature is Martin Feldstein (1974), which reports that, *ceteris paribus*, social security increases consumption substantially. As a result, the balanced growth paths for capital, output, and consumption are lowered substantially. Many other studies followed.[2] They reported mixed results, with perhaps a presumption that social security raises consumption *ceteris paribus*.

For the most part, this literature investigated the issue using regression models of the form

(1.1) $$c_t = b_0 + b_1 y_t + b_2 a_t + b_3 z_t + b_4' x_t + e_t ;$$

where c_t and y_t are per capita consumption and disposable income during period t; a_t and z_t are per capita private wealth and social security wealth at the beginning period t; x_t is a vector of other variables realized in period t; b_1, b_2 and b_3 are parameters; b_4 is a vector of parameters; and e_t is an error

[1] Some examples are Henry Aaron and Barry Bosworth (1997), Advisory Committee on Social Security (1997), David Altig and Jagadeesh Gokhale (1997), Peter Diamond (1996), Martin Feldstein (1996a), Feldstein and Andrew Samwick (1997), Edward Gramlich (1998), Laurence Kotlikoff and Jeffrey Sachs (1997), and Randall Mariger (1997).

[2] For example, see Robert Barro (1978); Barro and Glenn MacDonald (1979); Richard Burkhauser and John Turner (1982); Michael Darby (1978); Owen Evans (1983); Feldstein (1980, 1982, 1996); Feldstein and Anthony Pellechio (1979); Erkki Koskela and Matti Viren (1983); George Kopits and Padma Gotur (1980); Kotlikoff (1979); Dean Leimer and Selig Lesnoy (1982); Philip Meguire (1996); Franco Modigliani and Arlie Sterling (1979); Alicia Munnell (1974, 1980); and George von Furstenburg (1979). Ben Page (1998) exhaustively and critically reviews this literature.

term.[3] If b_3 turns out to be positive and statistically significant, social security is inferred to raise consumption *ceteris paribus*. Unfortunately, equation (1.1) is at best an approximate reduced form of the life-cycle model and at worst bears no relationship to any well-known structural model. If it can be interpreted as a good approximation to a reduced form, its parameters confound structural and expectational parameters. For example, even if Ricardian equivalence holds so that social security is completely neutral, the parameter b_3 can be positive if social security wealth is positively correlated with future disposable wage incomes, conditional on the other regressors. Conversely, even if b_3 is estimated to be insignificantly different from zero, social security may still increase consumption appreciably since social security wealth may be negatively correlated with future disposable wage incomes, conditional on the other regressors. In principle, these problems could be overcome by estimating a variant of equation (1.1) jointly with auxiliary forecasting equations for the components of future disposable wage income while imposing the cross-equation restrictions implied by rational expectations. In practice, this task is too formidable to attempt.

Another problem with equation (1.1) is that the variables included in it are likely to be difference stationary. If the error term e is also difference stationary, least squares is inconsistent. If instead it is mean stationary, the parameters are estimated superconsistently. Nevertheless, unless e is also uncorrelated with Δy, Δa, Δz, and Δx contemporaneously and at all leads and lags, the standard errors are inconsistent and standard procedures produce invalid inferences.[4] Moreover, correlation is required by the intertemporal budget constraint.

In this paper, I also attempt to estimate the effects of social security directly. My starting point is the aggregate Euler equation

(1.2) $$r_t = y_0 + y_1 \Delta \ln c_t + u_t ,$$

where r_t is the continuously compounded one-period real interest rate, ψ_1 is the relative risk aversion, and u_t is an error term with a zero mean and finite variance. Under Ricardian equivalence and some separability, aggregation and distributional assumptions, u_t is orthogonal to all lagged information. Consequently, if Ricardian equivalence holds and the augmented model

(1.3) $$r_t = y_0 + y_1 \Delta \ln c_t + y_2 (a_t + p_t + z_t) / c_t + u_t$$

[3] Some of the studies put per capita saving or per capita private assets in the left-hand member in lieu of per capita consumption. Clearly, such regressions can be rewritten in the form (1.1). Others do not deflate by population, a dubious but probably inconsequential practice.

[4] For proof of the previous three assertions and further discussion, see James Hamilton (1994), pp. 557-561, 586-589, 602-608.

is estimated using lagged instrumental variables, the estimate of ψ_2 should have a zero probability limit. In equation (1.3), π_t is disposable asset income in period t and u_t is an error term that has a zero mean and finite variance and is posited to be orthogonal to all lagged information. If instead ψ_2 is estimated to be significantly positive, fairly convincing evidence would then exist against either Ricardian equivalence or the other maintained assumptions.[5]

The intuition for a positive ψ_2 is straightforward. Suppose that Ricardian equivalence does not hold and, in particular, existing households do not completely offset additional government intergenerational transfers in their favor. Suppose further that households experience an unanticipated increase in wealth, which may take the form of budget deficits, upward revaluations of existing assets, windfall receipts of property income, or increases in the generosity of the social security system. This event results in a simultaneous jump in per capita consumption, followed by decreased growth in per capita consumption over some period of time in order to satisfy the aggregate intertemporal budget constraint. In a closed or large open economy, the initially higher and flatter path for per capita consumption would lower the balanced growth path for the capital stock and raise the balance growth path for the real interest rate.

This paper fits equation (1.3) to annual US data over the period 1950-1994, finding strong evidence that ψ_2 is positive and appreciable. It then quantifies the steady-state effects on the real interest rate of reforms in social security, defined in terms of how much they reduce the steady-state level of z_t/c_t. From these, it calculates the steady-state effects on the capital stock, output, and consumption.

The empirical work has four noteworthy implications. First, I am never able to reject the overidentifying restrictions of equation (1.3) even though those who have fitted variants of equation (1.2) typically reject them at low significance levels; e.g., Lars Peter Hansen and Kenneth Singleton (1983). Second, using the method advocated by John Campbell and Gregory Mankiw (1989, 1990, 1991), I find no evidence of binding constraints on borrowing in equation (1.3). Third, the parameter ψ_2 is highly significantly positive. As a result, Ricardian equivalence is decisively rejected. Fourth, social security reform is potentially capable of appreciably raising the balanced growth paths for the capital stock, output, and consumption.

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[5] My 1988 and 1993 papers and my joint 1994 paper with Iftekhar Hasan pursue a similar approach to investigate the effects of government debt on consumption. These papers omit social security wealth and property income from the Euler equation.

The remainder of the paper is organized as follows. Section II provides a theoretical discussion of why $(a_t+\pi_t+z_t)/c_t$ appears in model (1.3) if intergenerational connectedness is imperfect so that Ricardian equivalence fails. Section III reports the results from fitting models (1.2) and (1.3). Section IV investigates the effects of social security reform implied by the empirical model of section III. Finally, section V concludes.

II. THEORETICAL DISCUSSION

Suppose that households face perfect capital and insurance markets and maximize identical intertemporally and intratemporally separable objective functions characterized by a constant relative risk aversion and a constant subjective discount rate. In that case, the following first-order condition for a maximum holds:

(2.1) $$E_{t-1} \exp\left[r_t - r - a \ln(\hat{c}_t / \breve{c}_{t-1})\right] = 1,$$

where r_t is the continuously compounded real rate of return realized between periods *t*-1 and *t* on some financial asset, $\hat{c}_t$ and $\breve{c}_{t-1}$ are the per capita consumptions of the households that inhabit the economy in both periods *t* and *t*-1, ρ is the subjective discount rate, and α is the relative risk aversion. Given that the second and higher-order conditional moments of $r_t - a \ln(\hat{c}_t / \breve{c}_{t-1})$ are constant,[6] the Euler equation (2.1) can then be cast in the linear form[7]

(2.2) $$r_t = y_0 + a \ln(\hat{c}_t / \breve{c}_{t-1}) + u_t ,$$

where $y_0 \equiv r + E \ln E_{t-1} \exp[r_t - a \ln(\hat{c}_t / \breve{c}_{t-1})]$ and

(2.3) $$u_t \equiv (r_t - E_{t-1} r_t) - a[\ln(\hat{c}_t / \breve{c}_{t-1}) - E_{t-1} \ln(\hat{c}_t / \breve{c}_{t-1})] .$$

Equation (2.3) implies that the error term u_t is orthogonal to all information available in period *t*-1.

If everyone entering and exiting the economy over time is connected to existing households by operative altruistic bequest or gift motives, $\hat{c}_t$ and $\breve{c}_{t-1}$ can be measured by c_t and c_{t-1}, the per capita

[6] Christopher Carroll (1997) has argued that the variability of these moments is appreciable and persistent, implying that the error term *u* below is not orthogonal to all lagged information. An alternative interpretation of the results of this paper is therefore that these moments are positively and appreciably correlated with the ratio of the sum of private wealth, asset income, and social security wealth to consumption.

[7] If $\ln(\hat{c}_t / \breve{c}_{t-1})$ were directly observable, one could avoid making this distributional assumption since equation (2.1) could be directly estimated using the generalized method of moments. Unfortunately, imperfect proxies would have to be used, thereby introducing measurement error. The generalized method of moments does not produce consistent estimates in nonlinear models when the regressors are measured with error. Linearizing equation (2.1) enables this problem to be avoided.

consumptions in periods t and t-1. In that case, equation (2.2) can be implemented by applying the generalized method of moments (GMM) to

$$(2.4) \qquad r_t = \psi_0 + \psi_1 \Delta \ln c_t + u_t .$$

If only instrumental variables dated t-1 or earlier are used, the resulting estimate of ψ_1 should converge in probability to the relative risk aversion α. Furthermore, if equation (2.4) is modified to include additional variables, the coefficients on these variables should converge in probability to zero. Equation (2.4) is therefore eminently testable.

If disconnected households flow into or out of the economy, equation (2.4) does not hold. The reason is that the households exiting in period t-1 and entering in period t differ systematically from the households that inhabit the economy in both periods, leading $\Delta \ln c_t$ to differ systematically from $\ln(\hat{c}_t / \breve{c}_{t-1})$. For example, if consumption rises with age, one would expect the exiting households to have higher consumptions and the entering households to have lower consumptions than the population that spans both periods. In this case, their entry and exit pull the growth rate of per capita consumption downward relative to $\ln(\hat{c}_t / \breve{c}_{t-1})$. This effect tends to be larger, the greater are the entry and exit rates of disconnected households and the proportional consumption gaps.

Appendix A employs the demographic assumptions of Olivier Blanchard (1984) and Philippe Weil (1987) in order to derive an approximate expression for $\ln(\hat{c}_t / \breve{c}_{t-1})$. Specifically, all households are assumed to be disconnected from each other and to have a constant probability of exiting from the economy each period, and the number entering the economy each period is a constant fraction p of the population existing at the time.[8] In addition, it is assumed that the taxes levied to cover government consumption and to service the government debt are independent of age and that the wage rate does not vary systematically with age.[9] Although these assumptions lack realism, they do permit a tractable derivation of an Euler equation that can be implemented empirically. Despite their lack of realism, it may not be unreasonable to interpret the parameters of such an Euler equation structurally as is done in the next section.

On these assumptions,

$$(2.5) \qquad \ln(\hat{c}_t / \breve{c}_{t-1}) \approx \Delta \ln c_t + \left(\frac{p}{1-p}\right) m(a_t + \pi_t + z_t) / c_t ,$$

[8] The entry rate for the economy is best regarded as a metaphor for the extent to which altruistic intergenerational transfer motives are inoperative. On this interpretation, $1/p$ can be interpreted as the average number of periods before the representative dynasty encounters an inoperative bequest motive. In the limit as p approaches zero, intergenerational connectedness becomes perfect and the representative dynasty has an infinite horizon.

where μ be the mean of the marginal propensity to consume from wealth, a_t is per capita private wealth in period t, π_t is per capita asset income in period t, and z_t is the per capita present value of the current and future net transfers that existing households expect to receive as of period t less the per capita present value of the current and future net transfers that the newly arriving households expect to receive. I call this variable *social security wealth* because the social security system generates the lion's share of such redistributions and because the empirical analysis of the next section focuses on social security. The term in $(a_t+\pi_t+z_t)/c_t$ appears in equation (2.5) because the greater it is, the more existing households can consume relative to the newly arriving households, thereby depressing $\Delta\ln c_t$ relative to $\ln(\hat{c}_t / \breve{c}_{t-1})$. Substituting equations (2.5) into equation (2.2) yields

(2.6) $$r_t = \psi_0 + \psi_1 \Delta \ln c_t + \psi_2 (a_t + \pi_t + z_t) / c_t + u_t ,$$

where $\psi_2 \equiv \left(\frac{\pi}{1-\pi}\right)\alpha\mu$.

If GMM is applied to equation (2.6) with instrumental variables dated t-1 or earlier, the resulting estimates of ψ_1 and ψ_2 should converge in probability to α and $\left(\frac{\pi}{1-\pi}\right)\alpha\mu$, respectively. Therefore, the hypothesis of perfect intergenerational connectedness (p = 0) can be tested by examining whether the estimate of ψ_2 is statistically significant. A statistically insignificant estimate is consistent with the hypothesis, while a significantly positive estimate provides contrary evidence. Furthermore, the Euler equation (2.6) is just as readily testable as the standard Euler equation (2.4). One just modifies it to include additional variables of interest and tests whether the coefficients on these variables are significant. Under the assumptions underlying equation (2.6), one should not often reject zero coefficients for the additional variables.

Because social security wealth is not directly observable, one must find an adequate proxy for it. I employ a proxy based on the assumptions made above. In particular, social security is assumed to take the form of lump-sum taxes of $s_t/(1-\eta)$ levied on every household younger than age m and transfer payments of s_t/η given to every household at least m periods of age, where s_t is per capita expenditures on social security and η is the fraction of the population at least m periods of age. Using these assumptions, Appendix A shows that to a first approximation,

(2.7) $$z_t / c_t = \tfrac{1}{1-\eta} \sum_{i=0}^{m-1} (1+\delta)^{-i} E_t (s_{t+i} / c_{t+i}) ,$$

[9] An alternative, observationally equivalent assumption is that any systematic variation with age creates a difference in human capital proportional to consumption.

where $1+\delta$ is the mean of $\exp(r_t - \Delta \ln c_t)$. A proxy for z_t/c_t can then be obtained by developing a forecasting model for s/c and plugging its forecasts of $\{s_{t+i}/c_{t+i}\}_{i=0}^{m-1}$ into equation (2.7).[10]

III. Empirics

This section reports the results of fitting equation (2.4) and a variant of equation (2.6) to annual US data. The ultimate goal is to determine how much, if any, capital is crowded out by social security.

My measure for r_t is the difference between .7 times the average continuously compounded nominal interest rate on Treasury one-year discount bonds during the fourth quarter of year *t*-1 and the change in the logarithm of the deflator for expenditure on nondurable goods and services from the fourth quarter of year *t*-1 to fourth quarter of year *t*. This calculation assumes that the marginal tax rate on one-year Treasury discount bonds is .3. The results, however, are qualitatively similar for other plausible values of this marginal tax rate. For c_t, I use real expenditure on nondurable goods and services and real disposable wage income in the fourth quarter of year *t* divided by the mid-quarter resident population. The series a_t is real per capita private wealth at the end of year *t*-1 divided by the mid-quarter resident population for fourth quarter of year *t*-1. The series s_t is annual nominal old-age, survivors, disability, and health insurance benefits (OASDHI) plus public assistance to the old and disabled.[11] Finally, π_t is annual nominal property income during year *t* divided by the midyear resident population and the annual deflator for expenditures on nondurable goods and services. The data on $\Delta \ln c$, $\Delta \ln y$, and r span calendar years in order to avoid the statistical problems that result when overlapping averages are used in estimating Euler equations (Hall, 1988). Annual data are employed in the empirical analysis because the data on private wealth are not available quarterly. Appendix B provides a more complete description of the data and lists their sources.

Figure 1 shows that the path followed by $\ln(s_t/c_t)$ during the postwar period resembles the transition dynamics of a covariance-stationary process starting from an initial value far below its mean.

[10] One could instead use the proxy developed by Feldstein (1974) and updated by Feldstein (1996b). For the purposes of this paper, however, it is preferable to use a proxy based on the same assumptions as the theoretical model is. Otherwise, the estimated parameters may be difficult to interpret.

[11] Survivor, disability, and health insurance benefits are included because they are distributed to relatively old households and are financed by taxes on relatively young households. It is also important to include public assistance (SSI since 1974) to the old and disabled since they dominate OASDHI early in the sample period.

An adequate forecasting model for $x_t \equiv \ln(s_t/c_t)$ is the first-order autoregression[12]

$$(3.1)\qquad \hat{x}_t = \underset{(.0257)}{-.1050} + \underset{(.009)}{.944}\, x_{t-1}, \quad R^2 = .996, \quad SEE = .0430, \quad \text{Sample Period} = 1948\text{-}1995,$$

where the numbers in parentheses are standard errors. The autoregressive coefficient .944 differs statistically from one at any reasonable significance level since the Dickey-Fuller t-ratio is -6.30. Covariance-stationarity is thus supported. Equation (3.1) implies that the stationary distribution of s_t/c_t has a mean of .1531.[13] As a result, one can reasonably expect a further increase of about two percentage points in s_t/c_t from its value of .1311 in 1995 before it attains its unconditional mean.

The forecasts plugged into equation (2.7) take the form

$$(3.2)\qquad \{\exp[-1.885 + .944^i (x_t + 1.885)]\}_{i=0}^{m-1},$$

where -1.885 = -.1050/(1-.944) is the estimate of the mean of x_t implied by equation (3.1). Social security is assumed to be paid to 20 percent of households and to be supported by taxes levied on m = 45 cohorts of households. It is not at all clear what discount rate δ to use in equation (2.7). On the one hand, one might argue for a small value such as .02 or .03 per year since social security is purely redistributive, generating no aggregate risk. On the other hand, individual households are exposed to considerable political risk with regard to both benefits and taxes, not all of which may be diversifiable. Indeed, if $p > 0$, it cannot be entirely diversified away under any circumstances. In this case, one might argue for an appreciably larger discount rate. Here, I let the data speak for themselves by estimating the model

$$(3.3)\qquad r_t = \psi_0 + \psi_1 \Delta \ln c_t + \psi_2 (a_t + \pi_t)/c_t + \psi_3 (z_t/c_t)^* + u_t$$

for a range of discount rates, where $(z_t/c_t)^*$ is the proxy for z_t/c_t calculated as described above. I pay special attention to discount rates consistent with $\psi_2 = \psi_3$ since theory indicates that they should be equal. These discount rates turn out to be substantial, suggesting that households regard social security as quite risky.[14]

I used GMM to fit equation (2.4) to the data described above over the sample period 1951-1994. The instrumental variables included the intercept and two lags each of r_t, $\ln c_t$, $(a_t+\pi_t)/c_t$, and the $(z_t/c_t)^*$ for δ = .1182. (I explain why this is an appropriate value below.) The result was

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12 The Schwarz criterion selects this lag length from lag lengths ranging from one to four years. Although one might prefer a multivariate model in principle, finding variables with significant explanatory power for $\ln(s/c)$ is difficult.

13 In order to calculate this mean, I generated 1,000,000 observations from the fitted process with the estimated mean of x as the initial value and with normally distributed error terms. I then exponentiated them and calculated the sample mean of the resulting series.

$$\hat{r}_t = \underset{(.00342)}{.00392} - \underset{(.166)}{.041}\,\Delta\ln c_t\,, \quad SEE = .01645, \quad MSLJ = .0097. \tag{3.4}$$

The relative risk aversion is estimated to be negative, albeit insignificantly so. Furthermore, the marginal significance level of the *J*-statistic is low, indicating that the overidentifying restrictions of the model can be rejected at the .01 level. An important reason for the low value of *MSLJ* can be seen from examining the first eleven sample autocorrelations of the residuals, which are .61, .40, .33, .13, .04, -.04, -.25, -.31, -.37, -.41, and -.32. On the null hypothesis that the error term is serially uncorrelated, each of these should be approximately normally and independently distributed with a zero mean and a standard deviation of .16. This hypothesis is therefore roundly rejected.

Fitting equation (2.4) does not produce a reasonable statistical model for a straightforward reason. According to this equation, the real interest rate r_t is a linear combination of the white-noise process u_t and $\Delta\ln c_t$, which is not readily distinguishable from white noise. (Its first eleven sample autocorrelations are .18, -.16, .09, .10, -.18, -.12, .17, -.10, -.06, .11, and .07 with a large-sample standard error of .15.) The real interest rate should therefore be white noise itself. In fact, it is quite persistent since its first eleven sample autocorrelations are .47, .33, .35, .13, .05, -.03, -.24, -.28, -.32, -.35, and -.31. In order to obtain a reasonable statistical model, equation (2.4) should be modified to include some persistent regressors in addition to $\Delta\ln c_t$.

I used GMM to fit equation (3.3) to the data over the sample period 1951-1994. The instrumental variables included the intercept and two lags each of r_t, $\Delta\ln c_t$,, $(a_t+\pi_t)/c_t$, and $(z_t/c_t)^*$. Table 1 reports the results obtained for five discount rates ranging from .02 to .14. In all cases, the overidentifying restrictions easily pass muster, the estimates of ψ_2 and ψ_3 are highly significantly positive, and the estimates of ψ_1 are positive and border-line significant. The estimated Euler equations therefore provide strong evidence that $p > 0$. The relative risk aversion is estimated to be approximately .3. Furthermore, except for the estimate of ψ_3 and the marginal significance levels of the test statistic for the hypothesis $\psi_2 = \psi_3$, the results are basically identical for all of the discount rates considered. Thus, there is little basis on which to choose a discount rate other than whether $\psi_2 = \psi_3$ is consistent with the data. On that basis, the discount rate appears to be rather large, perhaps lying somewhere between .10 and .15. The differences in the estimates of ψ_3 observed in Table 1 are more apparent than real, however. The product of each estimate of ψ_3 and the sample standard deviation of

[14] Alternatively, they may be myopic.

the corresponding proxy for z/c ranges only from .0196 to .0204 for the five discount rates. As a result, $(z_t / c_t)^*$ induces basically the same sample variance in $\Delta \ln c_t$ whatever the discount rate δ might be.

The discount rate that equates the estimates of ψ_2 and ψ_3 is .1182. I used GMM to fit equation (3.3) with $\delta = .1182$ while imposing the restriction $\psi_2 = \psi_3$. I continued to use the same instrumental variables as for Table 1. The result was

(3.5) $$\hat{r}_t = \underset{(.045)}{-.357} + \underset{(.139)}{.295}\, \Delta \ln c_t + \underset{(.0075)}{.0598}\, [(a_t + p_t)/c_t + (z_t / c_t)^*],$$

$$SEE = .01207, \quad MSLJ = .6548, \quad \text{Sample Period} = 1950\text{-}1994.$$

The estimate of ψ_2 is highly significantly positive, and the estimate of ψ_1 is positive and statistically significant at the .05 level. Again, strong evidence is found for $p > 0$. The implied estimate of the relative risk aversion is approximately .3. It is thus less than many economists would expect it to be and fairly precisely estimated as well.[15] Finally, the first eleven sample autocorrelations of the residuals are .22, .09, .06, .03, -.01, -.06, -.27, -.16, -.19, -.39, and -.15. There is some hint of negative correlation at long lags, suggesting that it might be possible to improve on model (3.5). Nevertheless, the model does markedly improve on model (3.4).

The theory of the previous section assumed that households do not face binding constraints on borrowing. Campbell and Mankiw (1989, 1990, 1991) have argued that one can allow for the possibility of borrowing-constrained households by including the growth rate of per capita disposable wage income as an additional regressor in Euler equations.[16] For example, in the context of this paper, one would estimate

(3.6) $$r_t = y_0 + y_1 \Delta \ln c_t + y_2 [(a_t + p_t)/c_t + (z_t / c_t)^*] + y_4 \Delta \ln w_t + u_t$$

in lieu of equation (3.3), where w_t is real disposable wage income in the fourth quarter of year t divided by the mid-quarter resident population. An appreciable and significantly negative estimate of ψ_4 would then indicate that binding constraints on borrowing have important effects on consumption and asset prices.

I used GMM to fit equation (3.6) to the data over the sample period 1950-1994. The instrumental variables included the intercept and two lags each of r_t, $\Delta \ln c_t$, $(a_t + \pi_t)/c_t$, $(z_t / c_t)^*$, and $\Delta \ln w_t$. I set the discount rate to .1182 and tested whether the restriction $\psi_2 = \psi_3$ is consistent with the

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15 For example, many contributors to the real business-cycle literature calibrate α as 1, 2, or 3; see Thomas Cooley (1995).

16 Actually, Campbell and Mankiw estimate regressions with the growth rate of total disposable income though the logic of their approach would call for disposable wage income since borrowing-constrained households do not hold assets and have asset income.

data. When I found that it easily passed muster, I imposed it and reestimated. (The marginal significance level was .8952.) The final result was

(3.7)
$$\hat{r}_t = \underset{(.043)}{-.349} + \underset{(.159)}{.425}\,\Delta \ln c_t + \underset{(.0072)}{.0582}\,[(a_t + p_t)/c_t + (z_t/c_t)^*] - \underset{(.042)}{.021}\,\Delta \ln w_t,$$

$$SEE = .01180, \quad MSLJ = .3907.$$

The estimate of ψ_4 is precisely estimated but nonetheless statistically insignificant at any reasonable level.[17] Therefore, no evidence is found that constraints on borrowing have empirically important effects on aggregate consumption or asset prices.

IV. EFFECTS OF SOCIAL SECURITY REFORM

In this section, I calculate the steady-state values implied by a simple nonstochastic computational general-equilibrium model that embeds the estimate of equation (3.5) reported in Table 3. My purpose is to evaluate the economic significance of the estimates obtained in the previous section. The model is highly stylized and calibrated to match only the grossest features of the US economy. For this reason, the results are primarily useful in assessing economic significance, though they may give some idea of the size of social security's likely steady-state effects.

I assume that the economy is closed and has the intensive Cobb-Douglas production function

(4.1)
$$\tilde{q} = A\tilde{k}^{1/3},$$

where $\tilde{q}$ and $\tilde{k}$ are steady-state gross output and capital per efficiency unit of labor. I choose the parameter A so that the capital-output ratio is 2.5 years in the initial steady state in which $\tilde{q}$ is normalized to one per year. The choice of $1/3$ for the elasticity of output with respect to the capital stock is entirely conventional; e.g., Campbell (1994).

I assume that the government engages in only three activities: making purely exhaustive expenditures on goods and services; distributing social security benefits to the old; and servicing its debt. It finances these activities with a uniform marginal tax rate of .3 on the income from capital and a nondistorting tax on wage income.[18] I further assume that in the steady state, government expenditures on goods and services are 18 percent of gross output and the ratio of government debt to total

[17] Similar results hold for each of the discount rates considered in Table 1.

[18] One of the many benefits claimed for reforming social security is the reduction in tax distortions in the labor market. As pointed out by Murphy and Welch (1998), however, these benefits can be obtained in principle without fundamentally changing the character of social security. I therefore abstract from them here.

consumption is .653 years, roughly their values in the late 1990s. Finally, I assume that in the initial steady state, the ratio of social security wealth to total consumption is 1.31 years, an estimate of the unconditional mean ratio of social security wealth to total consumption.[19]

On average, capital depreciates .04 per year, the total replacement rate is .065 per year, and capital is paid its after-tax marginal product. The former figure is implied by the assumed capital-output rate of 2.5 years and the observation that depreciation in the US national income accounts is roughly 10 percent of GDP. The replacement rate is then consistent with a growth rate of approximately .025 per year in the number of efficiency units of labor as well as in the capital stock, gross output, and consumption. It then follows that

$$\tilde{r} = \ln\left(1 + (1-.3)\left[\tfrac{1}{3}(\tilde{q}/\tilde{k}) - .04\right]\right), \tag{4.2}$$

and

$$\tilde{c} = .82\tilde{q} - .065\tilde{k}\,, \tag{4.3}$$

where $\tilde{r}$ is the steady-state interest rate and $\tilde{c}$ is steady-state consumption per efficiency unit of labor.

I assume that per capita consumption and disposable wage income grow at a constant and common exogenous rate in the steady state. On this assumption, the steady-state real interest rate is given by

$$r = \kappa + .0661(\tilde{k}/\tilde{c})\exp(r) + .0661[.653\exp(r) + (z/c)]\,, \tag{4.4}$$

where the parameter κ is chosen to match the real interest rate implied by equations (4.1)-(4.3) in the initial steady state and the figure .0661 was obtained by multiplying the estimate .0582 in equation (3.6) by 1.105, the average ratio of total real expenditures on consumption goods to real expenditures on nondurable goods and services between 1950 and 1994.

Solving this model for the initial steady state yields

$$\tilde{k} = 2.50000$$
$$\tilde{q} = 1.00000$$
$$r = 0.06329$$
$$\tilde{c} = 0.65750$$

19 It is calculated as (1.1182/.1182)(.1531/1.105). The quantity 1.1182/.1182 is the capitalization factor implied by equation (2.7) for a discount factor of .1182; .1531 is the estimated unconditional mean of social security expenditures divided by consumer expenditures on nondurable goods and services; and 1.105 is the sample mean of the real consumption expenditure on nondurable goods and services divided by total real consumption expenditure over the period 1950-1994.

Now, consider the reform of social security. For the purposes of this paper, the reform is taken to be a reduction in the steady-state ratio of social security wealth to consumption with no offsetting increase in the steady-state ratio of government debt to consumption. Although many of the proposals put forward do entail temporary increases in the debt ratio and in practice might entail permanent increases, any such proposal has effects identical to those of a proposal that reduces social security wealth less in the first place with no offset. I therefore consider reform proposals that differ only in the ultimate extent to which they reduce the steady-state value of z/c.

Figure 2 plots the percentage increase in the steady-state capital stock per efficiency unit of labor against the percentage reduction in the steady-state ratio of social security wealth to consumption effected by the social security reform. The percentage increase in $\tilde{k}$ is essentially linear in z/c, rising about half a percent for each percent that z/c is reduced. Equation (4.1) then implies that $\tilde{q}$ increases by about one-sixth percent for each percent that z/c is reduced. As a result, any substantial privatization would appreciably increase the capital stock and gross output along their balanced growth paths.

Figure 3 plots the percentage increases in steady-state consumption per efficiency unit of labor that the increases in the capital stock plotted in Figure 2 would support. The curve is concave as one would expect from the near linearity of the curve in Figure 1 and is upward sloping, indicating that the US capital stock is well below its golden-rule level. A reform that completely eliminated social security wealth would raise consumption by about six percent in the steady state, a sizeable effect.

How do these effects compare with those obtained from computational general-equilibrium models? Kottlikoff, Smetters, and Walliser (1998) simulated the effects of the complete elimination of the old-age and survivors components of OASDHI, finding that such a change would raise the capital stock by 38.6 percent in the steady state of their model. These two components constitute about two-thirds of social security expenditure and presumably result in about two-thirds of social security wealth. Figure 2 implies an effect of about 36.3 percent in the model of this paper. This effect is remarkably close to what Kottlikoff, Smetters, and Walliser obtained. This finding suggests that the extensive detail in large-scale computational general-equilibrium models may not be necessary for determining the effects of policy changes on the aggregate economy.

V. CONCLUSIONS

This paper found strong empirical evidence that social security strongly affects the US economy. In particular, a decrease in the ratio of social security wealth to consumption initially lowers consumption and ultimately lowers the balanced growth path for the real interest rate and raises the balanced growth paths for the capital stock, output, and consumption. These effects are significant not only statistically but also economically. For each percent that a social security reform ultimately reduces the ratio of social security wealth to consumption, the balanced growth paths for the capital stock and output shift upward by about $\frac{1}{2}$ and $\frac{1}{6}$ percent, respectively. Moreover, a social security reform that eventually eliminates all social security wealth shifts the balanced growth path for consumption upward by about six percent.

The empirical investigation yielded in three interesting byproducts. First, unlike much of the previous literature, the Euler equation fitted here does not reject its overidentifying restrictions. Therefore, the addition of a term in wealth plus asset income to the standard Euler equation helps to fix it. Second, no evidence is found for binding constraints on borrowing once this term is included in the Euler equation. Third, Ricardian equivalence is resoundingly rejected.

APPENDIX A

For simplicity, I assume that both the exit and entrance rates are p. On that assumption, the fraction of the population that is age h is $p(1\text{-}p)^h$ for $h = 0, 1, 2, \cdots$. Let c_{ht} be the consumption of each household of age h in period t. The assumption of a constant exit probability implies that among unconstrained households, those exiting in period t-1 consume on average the same amount as those that survive from period t-1 to period t. Hence,

(A.1) $$\breve{c}_{t-1} = c_{t-1}.$$

Moreover, those that survive consist of all of the households existing in period t except the fraction p that have newly arrived. As a result,

(A.2) $$\widehat{c}_t = (c_t - pc_{0t}) / (1 - p).$$

From equations (A.1) and (A.2), it then follows that

(A.3) $$\ln(\widehat{c}_t / \breve{c}_{t-1}) = \ln[(c_t - pc_{0t}) / (1 - p)] - \ln c_{t-1}$$

$$= \Delta \ln c_t + \ln\left[1 + \left(\frac{p}{1-p}\right)\left(\frac{c_t - c_{0t}}{c_t}\right)\right]$$

$$\approx \Delta \ln c_t + \left(\frac{p}{1-p}\right)\left(\frac{c_t - c_{0t}}{c_t}\right)$$

Let w_{ht}, a_{ht}, and π_{ht} be the consumption, disposable wage income, beginning-of-period assets, and asset income of the unconstrained households of age h in period t. Because they completely annuitize their assets in perfect insurance markets and face perfect capital markets, their budget constraints take the form

(A.4) $$c_{ht} + a_{h+1,t+1} = w_{ht} + a_{ht} \exp(r_t) / (1-p)$$
$$= w_{ht} + (a_{ht} + \pi_{ht}) / (1-p), \quad h = 0, 1, 2, \cdots.$$

Solving equation (A.4) forward then yields

(A.5) $$c_{ht} + \sum_{i=1}^{\infty} (1-p)^i \exp\left(-\sum_{j=1}^{i} r_{t+j}\right) c_{h+i,t+i} = (a_{ht} + \pi_{ht}) / (1-p)$$
$$+ w_{ht} + \sum_{i=1}^{\infty} (1-p)^i \exp\left(-\sum_{j=1}^{i} r_{t+j}\right) w_{h+i,t+i}.$$

From equation (A.5) it follows that

(A.6) $$c_{ht} = \mu_{ht}\left[(a_{ht} + \pi_{ht}) / (1-p) + w_{ht} + \sum_{i=1}^{\infty} (1-p)^i \exp\left(-\sum_{j=1}^{i} r_{t+j}\right) w_{h+i,t+i}\right]$$

where

(A.7) $$\mu_{ht} \equiv \left\{1 + \sum_{i=1}^{\infty} (1-p)^i \exp\left(\sum_{j=1}^{i} [\ln(c_{s+j,t+j} / c_{s+j-1,t+j-1}) - r_{t+j}]\right)\right\}^{-1}.$$

The assumptions leading up to equation (2.2) imply that equation (A.6) reduces to

(A.8) $$\mu_{ht} = \mu_t \equiv \left\{1 + \sum_{i=1}^{\infty} (1-p)^i \exp\left[(\sigma - 1)\sum_{j=1}^{i} r_{t+j}\right]\right\}^{-1}.$$

As a result, equation (A.6) can be approximated by

(A.9) $$c_{ht} = \mu\left[(a_{ht} + \pi_{ht}) / (1-p) + \sum_{i=0}^{\infty} \gamma^i w_{h+i,t+i}\right],$$

where μ is the mean of μ_t and γ is the mean of (1-p)exp(-r_t).

Equation (A.9) reduces to

(A.10) $$c_{0t} = m\sum_{i=0}^{\infty} g^{i} E_t w_{i,t+i}$$

for h = 0 since a_{0t} = π_{0t} = 0. (Households start life without assets and asset income.) The age-earnings profile is given by

(A.11a) $$w_{st} = w_t - s_t \,/\, (1-h)$$

for households of ages 0, 1, ⋯, m-1 and

(A.11b) $$w_{st} = w_t + s_t \,/\, h$$

for households of ages m, m+1, m+2, ⋯, where w_t is the cross-cohort average disposable wage income and $\eta \equiv (1\text{-}p)^m$. Substituting equations (A.11a) and (A.11b) into equation (A.10) gives

(A.12) $$c_{0t} = m\left[\sum_{i=0}^{\infty} g^{i} E_t w_{t+i} - \sum_{i=0}^{m-1} g^{i} E_t s_{t+i} /(1-h) + \sum_{i=m}^{\infty} g^{i} E_t s_{t+i} / h\right].$$

Aggregating both members of equation (A.9) and substituting from equations (A.11a) and (A.11b) then produces

(A.13) $$c_t = m\left\{\sum_{h=0}^{\infty} p(1-p)^h (a_{ht} + p_{ht})/(1-p) + \sum_{i=0}^{\infty} g^{i}\left[\sum_{h=0}^{\infty} p(1-p)^h E_t w_{h+i,t+i}\right]\right\}$$

$$= m\left\{g^{-1}\left[p\cdot 0 + (1-p)\sum_{j=1}^{\infty} p(1-p)^j (a_{jt} + p_{jt})\right]/(1-p) + \sum_{i=0}^{\infty} g^{i} E_t w_{t+i}\right\}$$

$$+\sum_{h=m}^{\infty} p(1-p)^h \sum_{i=0}^{\infty} g^{i} E_t s_{t+i} / h + m\sum_{h=0}^{m-1} p(1-p)^h\left[-\sum_{i=0}^{m-1-h} g^{i} E_t s_{t+i} /(1-h) + \sum_{i=m-h}^{\infty} g^{i} E_t s_{t+i} / h\right]$$

$$= m\left\{a_t + p_t + \sum_{i=0}^{\infty} g^{i} E_t w_{t+i} + \sum_{i=0}^{\infty} g^{i} E_t s_{t+i}\right\}$$

$$+ m\left\{-\sum_{i=0}^{m-1}\left[p\sum_{j=0}^{m-1-i}(1-p)^j\right] g^{i} E_t s_{t+i} /(1-h) + \sum_{i=1}^{m-1}\left[p\sum_{j=1}^{i}(1-p)^{m-j}\right] g^{i} E_t s_{t+1} / h\right\}$$

$$+ m\left[p\sum_{h=0}^{m-1}(1-p)^h\right]\sum_{i=m}^{\infty} g^{i} E_t s_{t+i} / h \;= m\left\{a_t + p_t + \sum_{i=0}^{\infty} g^{i} E_t w_{t+1} + \sum_{i=0}^{\infty} g^{i} E_t s_{t+i}\right\}$$

$$+ m\left\{-\sum_{i=0}^{m-1}[1-(1-p)^{m-i}] g^{i} E_t s_{t+i} /(1-h) + (1-p)^m \sum_{i=1}^{m-1}[(1-p)^{-i} - 1] g^{i} E_t s_{t+1} / h\right\}$$

$$+ m[1-(1-p)^m]\sum_{i=m}^{\infty} g^{i} E_t s_{t+i} / h \;= m\left\{a_t + p_t + \sum_{i=0}^{\infty} g^{i} E_t w_{t+1} + \sum_{i=0}^{\infty} g^{i} E_t s_{t+i}\right\}$$

$$+m\left\{-\tfrac{1}{1-h}\sum_{i=0}^{m-1}g^{i}E_t s_{t+i}+\tfrac{h}{1-h}\sum_{i=0}^{m-1}R^{-i}E_t s_{t+i}+\sum_{i=1}^{m-1}R^{-i}E_t s_{t+1}-\sum_{i=1}^{m-1}g^{i}E_t s_{t+i}+\tfrac{1-h}{h}\sum_{i=m}^{\infty}g^{i}E_t s_{t+i}\right\}$$

$$=m\left\{a_t+p_t+\sum_{i=0}^{\infty}g^{i}E_t w_{t+1}-\tfrac{1}{1-h}\sum_{i=0}^{m-1}g^{i}E_t s_{t+i}+\tfrac{1}{h}\sum_{i=m}^{\infty}g^{i}E_t s_{t+i}+\tfrac{1}{1-h}\sum_{i=0}^{m-1}R^{-i}E_t s_{t+i}\right\},$$

where $1/R$ is the mean of exp($-r_t$). As a result,

(A.14)
$$\frac{c_t-c_{0t}}{c_t}=m\left[a_t+p_t)/c_t+\tfrac{1}{1-h}\sum_{i=0}^{m-1}R^{-i}(E_t s_{t+i}/c_t)\right].$$

$$=m\left\{(a_t+p_t)/c_t+\tfrac{1}{1-h}\left[s_t/c_t+\sum_{i=1}^{\infty}R^{-i}\exp\left(\sum_{j=1}^{i}\Delta\ln c_{t+j}\right)E_t\left(s_{t+i}/c_{t+i}\right)\right]\right\}$$

$$\approx m(a_t+p_t+z_t)/c_t,$$

where $1+\delta$ is the mean of exp($\Delta\ln c_t-r_t$) and z_t/c_t takes the form (2.7). Finally, substituting (A.14) into (A.3) generates equation (2.5).

APPENDIX B

Real expenditure on nondurable goods and services is the sum of lines 4 and 5 of NIPA Table 1.2. For 1947-1958, the fourth-quarter continuously compounded nominal interest rate on Treasury one-year discount bonds (*IQ4*) is the average of the September, October, November, and December values of the continuously compounded one-year interest rate on discount Treasury securities compiled by Huston McCulloch and Heon-Chul Kwon (1993). Using the 12-month Treasury bill rate reported in the Federal Reserve Bulletin on a discount basis, I calculated continuously compounded interest rates for October, November, and December of 1959-1994 and averaged them together to obtain *IQ4* for those years. Note that the data of McCulloch and Kwon are end-of-month observations so that averaging over the September, October, November, and December observations approximates an average over the fourth quarter. By contrast, the data from the Federal Reserve are monthly averages of daily figures. Real private wealth comes from Table B-115 of the 1995 *Economic Report of the President*; end-of-year figures for year *t*-1 are relabeled as beginning-of-year figures for year *t*. Real disposable income is nominal disposable income from line 1 of NIPA Table 2.1 divided by the deflator for expenditures on nondurable goods and services. The data on OASDHI and population come from lines 16, 19, and 35 of NIPA Table 2.1. Public assistance to the old and disabled comes from four annual statistical supplements to the *Social Security Bulletin*; viz., 1966, Table 113; 1976, Table 175; 1984-85, Table 175; 1996, Table 7.A4.) Beginning in 1974, this item also includes SSI for the aged and disabled, which replaced public assistance after 1975. Before-tax property income is national income minus compensation of employees minus a fraction ω of proprietors' income plus net government interest payments minus government interest payments to foreigners. These series are lines 1, 2, and 9 of NIPA Table 1.14 and lines 13 and 16 of NIPA Table 3.1. The fraction ω is the ratio of employee compensation to national income less proprietors' income. (The assumption being made here is that compensation for proprietors' labor services represents the same fraction of total proprietors' income as compensation is in the rest of national income.) After-tax asset income is before-tax asset income minus corporate profits taxes minus federal inheritance taxes minus other personal state and local taxes minus a fraction ζ of income taxes minus the inflation tax $-(1-CPI_{t-1}/CPI_t)\cdot GNFA_t$, where CPI_t is the consumer price index for December of year t and $GNFA_t$ is government net financial assets. These series are line 3 of NIPA Table 3.1, line 4 of NIPA Table 3.2, line 5 of Table 3.3, the sum of line 3 of NIPA Table 3.2 and line 3 of NIPA Table 3.3, page C-27 of the January/February issue of the *Survey of Current Business* and a printout provided by the Bureau of Economic Analysis, and Table B-114 of the 1995 *Economic Report of the President*. The fraction ζ is the ratio before-tax property income minus corporate profits taxes to national income plus net government interest payments minus corporate profits taxes. (The assumptions being made here are that property income is taxed at the same average rate as compensation and that only corporate profits taxes are deductible against income taxes.) All NIPA data come from a diskette provided by the Bureau of Economic Analysis.

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Table 1. Estimates of Equation (3.3) for Five Discount Rates

δ	***Estimate of ψ_1***	***Estimate of ψ_2***	***Estimate of ψ_3***	***SEE* [*MSLJ*]**	***MSL* $\psi_2=\psi_3$**
.02	.234 (.144) [.1035]	.0573 (.0070) [.0000]	.0221 (.0034) [.0000]	.01229 [.5311]	.0000
.05	.258 (.147) [.0790]	.0583 (.0072) [.0000]	.0328 (.0052) [.0000]	.01225 [.5316]	.0000
.08	.277 (.150) [.0645]	.0591 (.0075) [.0000]	.0444 (.0071) [.0000]	.01222 [.5298]	.0001
.11	.292 (.152) [.0557]	.0597 (.0076) [.0000]	.0565 (.0092) [.0000]	.01221 [.5271]	.4923
.14	.303 (.154) [.0501]	.0601 (.0077) [.0000]	.0686 (.0114) [.0000]	.01221 [.5240]	.1561

Notes. The sample period is 1950-1994. The numbers in parentheses and brackets below each estimate are its standard error and marginal significance level, respectively. *SEE* is the standard error of estimate; *MSLJ* is the marginal significance levels of the *J*-statistic, which tests the overidentifying restrictions. Column 6 reports the marginal significance level of the χ^2-statistic for testing the equality of ψ_2 and ψ_3.