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Modeling and Estimating the Effects of Institutional Variables on a Pay-as-you-go Social Security System and on Household Saving

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Abstract

Using a two-period overlapping generations model and three panel data sets of annual aggregate data from twenty-five countries, we estimate a fixed-effects Euler equation for household saving. We focus on the effects of several institutional and other variables, such as corruption and the debt to gross domestic product (GDP) ratio, on household saving and on the probability that a pay-as-you-go social security system will grant pensions. We find that social security contributions reduce saving in a less than one-for-one manner. Also, as corruption or the debt to GDP ratio increases, the probability that the system will grant pensions falls, and so does the effect of social security contributions on saving. Finally, the marginal effect of an improvement in the quality of institutions on the credibility of the social

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security system is greater in countries where the quality of institutions is low than in countries where it is high, a result that stresses the role of institutions in reducing uncertainty about pensions.

Keywords
pay-as-you-go, saving, institutions, corruption, debt to GDP ratio

The quality of institutions has long been recognized as an important determinant of economic outcomes. The purpose of this article is to model and estimate the effects of institutional and other variables, such as corruption and the government debt to gross domestic product (GDP) ratio, on the probability that a pay-as-you-go (PAYG) social security system will grant pensions to the old at retirement. Through this channel, we examine the effects of these variables on the relationship between social security contributions and household saving. To our knowledge, these effects have not been studied in the literature. This is surprising, since the credibility of the social security system has been a global issue during the past few decades (Gern 2002), and since the effect of social security on private saving has been a fervently debated issue, as it may have serious implications for capital formation and future income (see, e.g., Leimer and Lesnoy 1982, 606 and Congressional Budget Office [CBO] 1998, 1).

Many researchers consider the life cycle theory of consumption and saving a suitable framework for studying the effect of social security on household saving, which works through the intertemporal budget constraint. For example, if the present value of social security benefits exceeds the present value of social security taxes, lifetime resources increase, consumption increases in every period, and household saving falls (Kotlikoff 1979).

Feldstein (1974) argues that social security affects household saving in two ways. First, when the yield on social security taxes is equal to the market interest rate, social security benefits substitute fully for household saving. Second, when the retirement decision is endogenous, an increase in social security benefits may induce early retirement and hence a longer retirement period, which requires more saving during the working years. Social security benefits may also cause offsetting changes in private intergenerational transfers, thus reducing their depressing effect on household saving (Feldstein and Pellechio 1979, 362). Social security is a transfer from the young to the old. If benefits increase, parents may increase their bequests so as to offset the additional taxes paid by their children. The

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additional saving for these bequests offsets the reduction in saving caused by the additional taxes. Hubbard (1984) points out that, because of uncertainty over the length of life, even an actuarially fair fully funded system would reduce household saving by more than the social security tax.

Individuals may not always behave rationally or have perfect knowledge, however. Because of myopic behavior, some may fail to adjust their saving in response to changes in social security benefits (Feldstein and Pellechio 1979, 364). Also, according to Hubbard, Skinner, and Zeldes (1995), to the extent that social security ensures against uncertain events, it may reduce precautionary saving.

Clearly, theoretical considerations alone cannot determine the net effect of social security on household saving, hence the need for empirical research. This is stressed by the CBO (1998), which provides a useful summary of the empirical literature. In general, the empirical studies use standard consumption or saving functions but differ widely in model specification, data, and results. Here are some examples. First, cross-sectional studies using household data generally favor the idea that social security has a negative effect on household saving (Feldstein and Pellechio 1979; Kotlikoff 1979; King and Dicks-Mireaux 1982; Hubbard 1986; Bernheim 1987), although some studies find no significant effect (Gullason, Kolluri, and Panik 1993). Second, using cross-country data, where the cross-sectional units are countries and the observations are time averages, Barro and MacDonald (1979) find both positive and negative effects, Feldstein (1980) finds a negative effect, and Koskela and Virén (1983) and Graham (1987) find no effect. Third, using panel micro data, some studies find a negative effect (Attanasio and Rohwedder 2003). Finally, using time series data, Feldstein (1974, 1982) finds that social security reduces personal saving, but Leimer and Lesnoy (1982) question this finding.

Political and efficiency theories of social security consider political institutions as well as economic and demographic variables as determinants of social security spending. Based on voting models, political theories do not expect social security to emerge and grow in nondemocracies and assume that social security spending is higher in democracies. On the other hand, efficiency theories suggest that economic and demographic factors are more important determinants of social security spending than political institutions (Mulligan, Gil, and Sala-i-Martin 2002, 2–6). These studies do not find clear evidence on the relationship between political institutions and social security, however. They limit their attention to the institution of voting and to the form of the political system (democratic vs. nondemocratic), and neglect other variables, such as corruption and the debt to GDP ratio.

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Sticking to the standard neoclassical paradigm, in this article we use a two-period overlapping generations model (OGM) and annual aggregate data from twenty-five countries to estimate a fixed-effects Euler equation for household saving and focus on the effects mentioned earlier. We find that corruption and the debt to GDP ratio affect negatively the probability that a PAYG system will grant pensions. We also find that social security contributions negatively influence household saving in a less than one-for-one manner and that corruption and the debt to GDP ratio weaken this effect. When we consider two subpanels, one containing eleven Organization for Economic Cooperation and Development (OECD) countries, in which the quality of institutions is, on average, higher than that in the second subpanel of fourteen countries, we find that a ceteris paribus improvement in the quality of institutions increases the credibility of the social security system by more in countries where the quality of institutions is low than in countries where it is high.

After describing the theoretical model in the second section, we present the data and some preliminary tests (third section), describe our econometric methodology (fourth section), present the empirical results (fifth section), and conclude (sixth section).

The Model

Individual Behavior

We use a standard discrete-time two-period OGM. In period t, there are L_t young and L_{t-1} old individuals. Each young individual offers inelastically one unit of labor in period t and receives a real wage (w_t), which he or she disposes for current consumption (c_{1t}), social security contributions (d_t), and future consumption (c_{2t+1}), which requires positive current saving ($s_t = w_t - c_{1t} - d_t > 0$). In period $t + 1$, the retirement period, he or she finances consumption using his or her savings (including interest) and the social security benefits (b_{t+1}), leaving no bequests. We assume that each individual represents a household and that employment grows at an exogenous rate n, that is, $L_t = (1 + n)L_{t-1}$.

Following the standard practice, we use a constant relative risk aversion utility function of a young individual, so his or her lifetime utility is:[1]

$$U = \frac{c_{1t}^{1-\gamma}}{1-\gamma} + \frac{1}{1+\rho} E_t\left(\frac{c_{2t+1}^{1-\gamma}}{1-\gamma}\right), \gamma > 0, \frac{1}{1+\rho} > 0, \quad (1)$$

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where ρ is the rate of time preference, γ is the coefficient of relative risk aversion, and E_t is the rational expectations operator conditional on information up to time t.

The intertemporal budget constraint of the individual is given by:

$$E_t(c_{2t+1}) = (1+r_t)s_t + E_t(b_{t+1}), \tag{2}$$

where we assume that the real interest rate (r_t) is associated with an inflation-protected bond, so it is known at the beginning of period t, when the decision about consumption and saving for t is made.[2] The Euler equation for consumption is $E_t(c_{1t}/c_{2t+1})^{\gamma} = (1+\rho)/(1+r_t)$, hence the Euler equation for saving (s_t) is (see Appendix A that is available upon request):

$$\frac{(1+r_t)s_t + E_t(b_{t+1})}{w_t - s_t - d_t}\left(\frac{1+\rho}{1+r_t}\right)^{1/\gamma} = E_t\left[\frac{1}{(1+e_{t+1})^{1/\gamma}}\right], \tag{3}$$

where e_{t+1} is a rational expectations error.

Introducing a PAYG Social Security System

Under a PAYG system, the social security contributions (d_t) paid by the young individuals finance the benefits (b_t) paid to the old in the same period. Since $L_t/L_{t-1} = 1+n$, that is, to each old individual there correspond $1+n$ young individuals, it follows that the contributions finance an amount of benefits $(1+n)d_t$ per person. Thus, letting $p(\boldsymbol{x}_t)$ denote the probability that the social security system will grant pensions, the expected benefits of period $t+1$ can be defined as:[3]

$$E_t(b_{t+1}) = p(\boldsymbol{x}_t)(1+n)d_{t+1}. \tag{4}$$

The $k \times 1$ vector $\boldsymbol{x}_t$ contains variables like an index of corruption (IC_t), with higher values indicating *lower* corruption, the government debt to GDP ratio (DG_t), and so on. We assume that the better the quality of institutions and the lower the debt to GDP ratio, the greater the probability that the system will grant pensions, for example, $\partial p(\boldsymbol{x}_t)/\partial(IC_t) > 0, \partial p(\boldsymbol{x}_t)/\partial(DG_t) < 0$, and so on, where the probability $p(\boldsymbol{x}_t)$ is determined by a logit model,[4] that is, $p(\boldsymbol{x}_t) = 1/\left(1+e^{-\beta' x_t}\right)$. We are interested in the effects of DG_t, IC_t, and so on, on the derivative $\partial s_t/\partial d_t$. Substituting Equation (4) into Equation (3) and rearranging yields:

$$(1+r_t)s_t + p(\boldsymbol{x}_t)(1+n)d_{t+1} = [(1+r_t)/(1+\rho)]^{1/\gamma}(w_t - s_t - d_t)u_{t+1}, \tag{5}$$

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where $u_{t+1} = E_t\left[1\Big/(1+e_{t+1})^{1/\gamma}\right]$. We now calculate the derivative $\partial s_t/\partial d_t$ by applying the implicit function theorem to Equation (5). For simplicity, we set e_{t+1} equal to its expected value, which is zero, so $u_{t+1} = 1$.[5] Also, we set $d_{t+1} = (1 + r_d)d_t$, where r_d is the growth rate of d_t, assumed to be constant. The result is:

$$\frac{\partial s_t}{\partial d_t} = -\frac{p(\boldsymbol{x}_t)(1+n)(1+r_d) + [(1+r_t)/(1+\rho)]^{1/\gamma}}{1 + r_t + [(1+r_t)/(1+\rho)]^{1/\gamma}} < 0. \qquad (6)$$

Thus, the probability $p(\boldsymbol{x}_t)$ reinforces the negative effect of social security contributions on household saving. For example, as the value of the index IC_t increases (less corruption), $p(\boldsymbol{x}_t)$ increases, and the negative number $\partial s_t/\partial d_t$ becomes more negative, that is, it decreases, $\partial(\partial s_t/\partial d_t)/\partial(IC_t) < 0$. As another example, as DG_t increases, $p(\boldsymbol{x}_t)$ decreases, and the expected social security benefits also decrease (see Equation (4)). As a result, when social security contributions increase, rational individuals, who want to secure a certain level of consumption at retirement, will reduce their saving by less, in an attempt to self-insure themselves against the higher uncertainty induced by higher indebtedness. That is, as DG_t increases, the negative number $\partial s_t/\partial d_t$ will become less negative, that is, it will increase, $\partial(\partial s_t/\partial d_t)/\partial(DG_t) > 0$.

The Data

We use three panel data sets in our empirical analysis. The first is a balanced panel of eleven OECD countries, namely, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, the United Kingdom, and the United States, for which we have data for the time period 1984 to 2009. The second is also a balanced panel of fourteen different countries, namely, Austria, Cyprus, Czech Republic, Estonia, Greece, Hungary, Latvia, Lithuania, Norway, Poland, Portugal, Slovakia, Spain, and Sweden, for which we have data for the time period 1995 to 2009. The third panel is unbalanced and consists of all the above twenty-five countries. This selection of countries was dictated by data availability and the fact that these countries generally use a PAYG system.[6] The sources of the data are as follows: (1) the annual macroeconomic database of the European Commission (AMECO), (2) World Development Indicators (WDI), (3) International Financial Statistics (IFS), and (4) International Country Risk Guide (ICRG).

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The empirical definitions of the variables are as follows: s_t = household saving per employee (AMECO),[7] d_t = social security contributions per employee (AMECO), w_t = real wage constructed by deflating a nominal wage rate (AMECO), $gdpe_t$ = GDP per employee (AMECO), dbe_t = general government debt per employee (AMECO), dfe_t = general government deficit per employee (AMECO), $DG_t = dbe_t/gdpe_t$, $FG_t = dfe_t/gdpe_t$, TE_t = total employment (AMECO), n_t = growth rate of TE_t, TU_t = total unemployment (AMECO), nu_t = growth rate of TU_t, ER_t = exchange rate (number of units of national currency per euro, AMECO), P_t = consumer price index (CPI, 2005 = 100, WDI), PE_t = percentage change in CPI (WDI), r_t = ex-post real interest rate constructed by subtracting PE_t from an interest rate on Treasury Bills or certificates of deposit or interbank loans, each with three months of maturity (IFS, AMECO),[8] and GP_t = GDP deflator (2005 = 100, AMECO). The variables s_t, d_t, w_t, $gdpe_t$, dbe_t, and dfe_t are expressed in thousands of euros. The variables s_t, d_t, and w_t are deflated by P_t, whereas the variables dbe_t, dfe_t, and $gdpe_t$ are deflated by GP_t.

Regarding the institutional variables, we use some indices of political risk from the ICRG, namely, IC_t = an index of corruption, which takes on values between 0 and 6 (again, higher values of IC_t indicate less corruption); PS_t = an index of government stability (reflecting popular support, government unity, and legislative strength), which takes on values between 0 and 12, with higher values indicating more stability; SC_t = an index of socioeconomic conditions (reflecting unemployment, consumer confidence, and poverty), which also takes on values between 0 and 12, with higher values indicating better conditions; and DA_t = an index of democratic accountability (reflecting the responsiveness of the government to its people), which takes on values between 0 and 6, with higher values indicating more accountability. Thus, we assume that $\partial p(\boldsymbol{x}_t)/\partial(IC_t) > 0$, $\partial p(\boldsymbol{x}_t)/\partial(PS_t) > 0$, $\partial p(\boldsymbol{x}_t)/\partial(SC_t) > 0$, and $\partial p(\boldsymbol{x}_t)/\partial(DA_t) > 0$.

As always, "a prelude to further inference" is to test for unit roots (Maddala and Kim 1998, 139) to ensure that we avoid the spurious regression problem (Kao 1999). To this end, we use (1) the t^*-test of Levin, Lin, and Chu (2002, LLC), (2) the t test of Breitung (2001), where we include both an individual constant and a time trend, and (3) the Fisher-type augmented Dickey–Fuller (ADF) test. Table 1 reports the results produced by the computer program *EViews* 6. In each panel, for each variable, at least one test suggests stationarity, so we take all of them to be $I(0)$.

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Table 1. Panel Unit-root Tests.

Test	LLC		Breitung	Fisher–ADF		Decision
Variable	t_μ^*	t_τ^*	t	χ_μ^2	χ_τ^2	$I(1)$ or $I(0)$?
Part A. The eleven-country panel						
s_t	−2.8***	−0.4	0.4	46.6***	29.8	$I(0)$
d_t	−4.4***	−2.1**	0.0	48.9***	39.5**	$I(0)$
w_t	−1.6*	−1.1	−1.3	31.8*	32.3*	$I(0)$
r_t	−1.2	−2.7***	−3.3***	15.8	49.4***	$I(0)$
DG_t	1.2	0.1	−1.9**	11.6	21.4	$I(0)$
FG_t	−0.9	3.8	4.5	36.1**	18.4	$I(0)$
$gdpe_t$	−3.8***	4.9	1.0	23.3	12.9	$I(0)$
n_t	−1.5*	0.2	0.5	61.7***	47.6***	$I(0)$
nu_t	−1.2	1.0	2.0	65.4***	46.0***	$I(0)$
IC_t	−0.3	−0.1	−1.8**	10.7	23.4	$I(0)$
PS_t	−1.1	0.2	−3.2***	31.7*	21.1	$I(0)$
SC_t	−0.4	−0.3	−2.3***	23.7	19.0	$I(0)$
DA_t	−0.5	−0.1	−2.2**	19.8	14.3	$I(0)$
Part B. The fourteen-country panel						
s_t	−2.2**	0.3	3.5	53.8***	37.9*	$I(0)$
d_t	2.6	−2.9***	1.5	9.0	43.2**	$I(0)$
w_t	−2.5***	−0.8	4.5	26.5	33.9	$I(0)$
r_t	−6.6***	−3.9***	1.2	69.2***	54.8***	$I(0)$
DG_t	−1.8**	−0.9	3.9	29.2	36.3	$I(0)$
FG_t	−2.9***	3.5	5.6	64.2***	29.1	$I(0)$
$gdpe_t$	−5.3***	4.5	6.1	30.3	18.8	$I(0)$
n_t	0.9	2.1	5.7	40.7*	29.9	$I(0)$
nu_t	2.5	6.2	7.8	39.5*	28.5	$I(0)$
IC_t	−3.8***	−4.4***	−2.8***	29.9	37.9*	$I(0)$
PS_t	−6.4***	−6.0***	−0.1	67.9***	85.3***	$I(0)$
SC_t	−5.4***	−12.0***	0.3	46.9**	63.9***	$I(0)$
DA_t	−5.8***	−2.9***	1.4	42.7***	17.8	$I(0)$
Part C. The twenty-five-country panel						
s_t	−3.5***	0.5	2.2	103.0***	72.1**	$I(0)$
d_t	−1.5*	−3.1***	0.9	58.7	82.3***	$I(0)$
w_t	−2.3**	−1.2	3.4	49.1	79.2**	$I(0)$
r_t	−4.7***	−4.0***	−0.5	85.1***	97.9***	$I(0)$
DG_t	2.6	0.1	0.2	38.1	72.8**	$I(0)$
FG_t	−1.9**	5.1	7.1	99.8***	47.5	$I(0)$
$gdpe_t$	−5.9***	6.6	9.6	54.3	31.4	$I(0)$
n_t	0.3	1.8	4.4	97.3***	73.2**	$I(0)$
nu_t	1.6	5.6	6.6	100.0***	68.6**	$I(0)$

(continued)

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Table 1. (continued)

Test	LLC		Breitung	Fisher–ADF		Decision
Variable	t_μ^*	t_τ^*	t	χ_μ^2	χ_τ^2	$I(1)$ or $I(0)$?
IC_t	-2.2^{**}	-2.7^{***}	-3.1^{***}	40.7	61.3^{*}	$I(0)$
PS_t	-4.5^{***}	-3.2^{***}	-2.4^{***}	99.6^{***}	106.0^{***}	$I(0)$
SC_t	-4.1^{***}	-9.2^{***}	-1.3^{*}	70.7^{**}	82.9^{***}	$I(0)$
DA_t	-3.6^{***}	-2.4^{***}	-0.6	62.5^{***}	32.1	$I(0)$

Note: The subscripts μ and τ indicate the presence of individual constant and constant and trend, respectively. In each cross-section ADF regression, the lag length is chosen by the Schwartz criterion. In the LLC test, a kernel-based consistent estimator of the residual covariance is used. *EViews* 6 has produced these results. ADF = augmented Dickey–Fuller; LLC = Levin, Lin, and Chu.
***, **, and * indicate statistical significance at the 1, 5, and 10 percent level.

Econometric Methodology

Taking logarithms in Equation (5), substituting $p(\boldsymbol{x}_t) = 1/\left(1 + e^{-\boldsymbol{\beta}'\boldsymbol{x}_t}\right)$, and adding country-specific dummies (D_i) yields our estimating equation:

$$\begin{aligned} &-\ln(1+\rho) + \Sigma_{i=1}^{N-1}\delta_i D_i + r_{it}^* + \gamma c_{it}^* \\ &\quad - \gamma \ln\left[(1+r_{it})s_{it} + \frac{1}{1+e^{-\boldsymbol{\beta}'\boldsymbol{x}_{it}}}(1+n_{it})d_{it+1}\right] = u_{it+1}^*, \end{aligned} \tag{7}$$

where $r_{it}^* = \ln(1 + r_{it})$, $c_{it}^* = \ln(w_{it} - s_{it} - d_{it})$, $u_{it}^* = -\gamma \ln u_{it}$, and $N =$ the number of countries.[9] The intercept in Equation (7), $-\ln(1+\rho) \approx -\rho$, corresponds to the reference country, for which the dummy (D_N) is omitted. Hence, the country-specific effect δ_i can be interpreted as the difference in the negative of the rate of time preference between country i and the reference country. The total number of parameters to be estimated is $K = N + k + 1$, that is, the two utility-function parameters (ρ and γ), the k parameters in the vector $\boldsymbol{\beta}$, and the $N - 1$ coefficients of the dummies (the δs).

We estimate Equation (7) by the generalized method of moments (GMM) using the computer program *WinRATS* 7.0 and six sets of instrumental variables (IVs). As in Campbell and Mankiw (1990, 268), the endogenous variables are lagged at least twice before they are used as IVs, so they are dated $t - 1$ or earlier, since the error term is dated $t + 1$. Thus, a variable dated $t - 1$ is a *two*-period lag in the computer program, a variable dated $t - 2$ is a *three*-period lag, and so on. A variable is included in an IV set if (1) it is statistically significant in the regression of each of the endogenous variables c_{it}*, s_{it}, and r_{it}* on the set of potential IVs and (2) it

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helps to achieve empirical identification (i.e., correct signs and statistical significance) of as many parameters as possible. Thus, in the case of the eleven-country panel, we use the following two IV sets: $V_1 = (1, D_1, \ldots, D_{10}, r^*_{it-2}, c^*_{it-2}, s_{it-1}, s_{it-2}, n_{it-3}, d_{it-2}, IC_{it}, SC_{it}, SC_{it-1}, DA_{it}, DA_{it-2}, DA_{it-3}, DG_{it-1}, DG_{it-2}, FG_{it-1}, FG_{it-2}, gdpe_{it-1}, gdpe_{it-2}, gdpe_{it-3}, nu_{it-2})$, which contains $M = 31$ IVs; and $V_2 = (1, D_1, \ldots, D_{10}, r^*_{it-2}, c^*_{it-2}, s_{it-1}, s_{it-2}, s_{it-3}, n_{it-2}, d_{it-1}, d_{it-2}, IC_{it}, IC_{it-1}, SC_{it}, SC_{it-1}, PS_{it}, PS_{it-1}, DG_{it-2}, FG_{it-2}, gdpe_{it-1}, gdpe_{it-2}, gdpe_{it-3}, nu_{it-1})$, where again $M = 31$.

In the case of the fourteen-country panel, we use the IV sets $V_1' = (1, D_1, \ldots, D_{13}, r^*_{it-1}, r^*_{it-2}, c^*_{it-1}, c^*_{it-2}, s_{it-1}, n_{it-2}, d_{it-1}, IC_{it}, IC_{it-1}, PS_{it}, PS_{it-1}, SC_{it}, SC_{it-2}, DA_{it}, DA_{it-2}, DG_{it-2}, FG_{it-1}, FG_{it-2}, gdpe_{it-1}, gdpe_{it-2}, nu_{it-1}, nu_{it-2})$ $(M = 36)$, and $V_2' = (1, D_1, \ldots, D_{13}, r^*_{it-1}, r^*_{it-2}, c^*_{it-1}, c^*_{it-2}, s_{it-1}, n_{it-1}, d_{it-1}, IC_{it}, IC_{it-2}, PS_{it}, PS_{it-1}, SC_{it}, SC_{it-1}, DA_{it}, DA_{it-2}, DG_{it-2}, FG_{it-1}, FG_{it-2}, gdpe_{it-1}, gdpe_{it-2}, nu_{it-1}, nu_{it-2})$ $(M = 36)$.

Finally, in the case of the twenty-five-country panel, we use the IV sets $V_1'' = (1, D_1, \ldots, D_{24}, r^*_{it-1}, r^*_{it-2}, c^*_{it-1}, c^*_{it-2}, s_{it-1}, n_{it-2}, d_{it-1}, d_{it-2}, IC_{it}, SC_{it}, SC_{it-1}, PS_{it}, PS_{it-2}, DA_{it}, DA_{it-2}, DG_{it-2}, FG_{it-2}, nu_{it-2})$ $(M = 43)$ and $V_2'' = (1, D_1, \ldots, D_{24}, r^*_{it-1}, r^*_{it-2}, c^*_{it-1}, c^*_{it-2}, s_{it-1}, s_{it-2}, n_{it-1}, d_{it-1}, IC_{it}, IC_{it-2}, SC_{it}, SC_{it-1}, PS_{it}, PS_{it-2}, DA_{it}, DA_{it-2}, DG_{it-1}, DG_{it-2}, FG_{it-1}, FG_{it-2})$ $(M = 45)$.

As is well known, in addition to being uncorrelated with the error term, the IVs must be correlated with the endogenous variables. If this correlation is weak, there arises the well-known weak-instrument problem, and the parameters are only weakly identified, meaning that biases are likely to occur (Stock, Wright, and Yogo 2002; Inoue and Rossi 2011).

A related issue is the number of IVs (M) and their lag lengths, see, for example, Tauchen (1986) for time series data and Ziliak (1997) for panel data. This literature demonstrates that, as M increases, GMM estimates tend to be biased downward, and this bias outweighs the gains in efficiency. Thus, Tauchen's (1986, 410, 415) advice is to keep the lag lengths of the IVs short. In general, the parameters are weakly identified when their estimates are sensitive to the addition of IVs or to changes in the sample (Stock, Wright, and Yogo 2002, 527).

In the case of the "difference GMM" linear regression, where the data are differenced in order to eliminate the fixed effects, Roodman (2009) argues that using too many IVs tends to bias GMM estimates in the direction of ordinary least squares (OLS) and to lead to implausibly high p values of the well-known J statistic for testing instrument validity (and model

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specification), which is asymptotically distributed as χ^2_{M-K}. He recognizes that there is no precise guidance on the choice of M and recommends that the ratio j/N should not be too high "in some vague sense" (p. 140, 142), where j is the number of IVs, excluding the dummies for the fixed effects, since the latter are differenced away, that is, $j = M - (N - 1)$. Roodman (2009, 151–53) criticizes two articles, where the ratio j/N is close to unity, by arguing that such values of j/N are high enough to cause the above problems. On the other hand, Ziliak (1997, 424–25) finds that as j/N increases the J test tends to over-reject, implying *low p* values of the J test. This result alleviates Roodman's concern regarding the reliability of the J test.

We address these issues, which can be subsumed under the more general term "weak identification," by performing the following tasks. First, we calculate the values of R^2 and the corresponding values of the standard F statistic from the OLS regressions of each of the endogenous variables (c_{it}*, s_{it}, and r_{it}*) on the IVs. We obtain high values of R^2 and values of F that satisfy the well-known "$F > 10$ rule," suggesting that our GMM regressions escape the weak-instrument problem (Stock, Wright, and Yogo 2002, 522). For example, in the case of the eleven-country panel, when V_2 is used, these values of R^2 are .95, .92, and .67, and the values of F are 120.8, 83.9, and 13.7 (p value $= 0$ in every case).

Second, we also produce nonlinear least squares (NLLS) estimates. The GMM estimates will be deemed reliable if they differ substantially from the NLLS ones.

Third, we conduct sensitivity analysis in the following two directions: (1) by changing the sample, namely, by using three panels, which differ from each other in the countries included and (one of them) in the sample period and (2) by dropping the IVs with the "deepest" lag, that is, those dated $t - 3$ (Tauchen 1986, 410), and reestimating Equation (7); if the coefficient estimates remain fairly stable, but their t-statistics become smaller (in absolute value), we will conclude that the IVs dated $t - 3$ add to efficiency without introducing much bias and will keep them.

Results

Tables 2 and 3 report the empirical results for each panel separately. Before considering them, note the following: (1) in every estimation discussed subsequently we use clustered standard errors, which allow for arbitrary patterns of serial correlation and heteroscedasticity, (2) the estimates seem to be robust to the choice of starting values for the parameters,[10] and (3) based on the theory of the second section, we view

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Table 2. GMM and NLLS Estimation of Equation (7).

	$\hat{\rho}$	$\hat{\gamma}$	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	J
Part A. The eleven-country panel						
GMM ($\boldsymbol{V}_1$)	.0978*** (7.36)	.0429*** (6.70)	−15.33** (−2.03)	1.78** (1.65)	−0.62 (−1.00)	0.019 (1.0)
GMM($\boldsymbol{V}_2$)	.0978*** (6.55)	.0436*** (5.80)	−10.82*** (−2.61)	0.73** (1.98)	—	0.027 (1.0)
NLLS	.0715*** (6.13)	.0298*** (5.08)	−10.87*** (−2.61)	0.67 (1.21)	—	—
Part B. The fourteen-country panel						
GMM($\boldsymbol{V}_1'$)	.0540*** (2.67)	.0359** (1.79)	−3.23** (−2.15)	3.55 (1.14)	−0.95 (−1.25)	0.026 (1.0)
GMM($\boldsymbol{V}_2'$)	.0716*** (4.96)	.0485*** (2.91)	−2.84* (−1.42)	0.63** (1.75)	—	0.028 (1.0)
NLLS	.0674*** (4.85)	.0448*** (2.65)	−1.56 (−0.90)	0.75 (0.66)	—	—
Part C. The twenty-five-country panel						
GMM($\boldsymbol{V}_1''$)	.0782*** (4.33)	.0371*** (3.79)	−4.55*** (−6.10)	0.61* (1.38)	−0.15 (−1.07)	0.054 (1.0)
GMM($\boldsymbol{V}_2''$)	.0845*** (6.59)	.0405*** (7.21)	−5.09*** (−3.27)	0.42 (1.14)	—	0.066 (1.0)
NLLS	.0612*** (6.16)	.0323*** (5.66)	−3.18** (−1.68)	0.51 (1.18)	—	—

Note: The values in parentheses following coefficient estimates are *t* statistics, while those following the values of the *J* statistic are *p* values. As a rule, we keep in the regression a variable whose coefficient has a *t* ratio larger than 1 in absolute value. These results have been produced by the econometric computer program *WinRATS* 7.0. GMM = generalized method of moments; NLLS = nonlinear least squares.
***, **, and * indicate statistical significance at the 1, 5, and 10 percent level, respectively, assuming a one-sided alternative hypothesis (see the second, third, and fifth sections).

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Table 3. Estimates of the Derivatives of Interest.

$\frac{\partial p(\bar{\mathbf{x}})}{\partial(DG)}$	$\frac{\partial p(\bar{\mathbf{x}})}{\partial(IC)}$	$\frac{\partial s}{\partial d}$	$\frac{\partial\left(\frac{\partial s}{\partial d}\right)}{\partial(DG)}$	$\frac{\partial\left(\frac{\partial s}{\partial d}\right)}{\partial(IC)}$
Part A. The eleven-country panel				
−.16 (−0.62)	.01 (0.54)	−.20*** (−6.50)	.13 (0.62)	−.01 (−0.54)
Part B. The fourteen-country panel				
−.52 (−0.69)	.12** (1.95)	−.83*** (−3.22)	.40 (0.67)	−.09** (−1.81)
Part C. The twenty-five-country panel				
−.91 (−1.27)	.08 (0.66)	−.38** (−1.84)	.73* (1.33)	−.06 (−0.68)

Note: These derivatives and their approximate standard errors are derived in Appendix B, which is available from the authors upon request. They have been evaluated at the sample means of the variables and at the parameter estimates from the preferred regressions. The standard errors have been calculated using the well-known formula (Kmenta 1971, 444). The *t* statistics are in parentheses following the estimates.
***, **, and * indicate statistical significance at the 1, 5, and 10 percent level, respectively, assuming a one-sided alternative hypothesis (see the second, third, and fifth sections).

all of our tests of significance on the parameters and the derivatives of interest as one sided.

The Eleven-country Panel

We begin with the results from the eleven-country panel. First, we estimate Equation (7) using the vectors $\boldsymbol{x}_{it} = (DG_{it}, IC_{it}, SC_{it})'$,[11] $\boldsymbol{\beta}' = (\beta_1, \beta_2, \beta_3)$, where $\beta_1 < 0$, $\beta_2 > 0$, and $\beta_3 > 0$, and $\boldsymbol{V}_1$ ($j/N = 1.9$). This is the first regression reported in Part A of table 2. All the coefficients are correctly signed and statistically significant at the 5 percent level, except for the estimate of β_3, which is not significant at conventional levels.

Thus, second, we exclude SC_{it} from the vector $\boldsymbol{x}_{it}$ and reestimate Equation (7), where now $\boldsymbol{\beta}' = (\beta_1, \beta_2)$, using the IV set $\boldsymbol{V}_2$ ($j/N = 1.9$). This is the second regression in Part A of table 2, where all coefficients have the expected sign and are significant at the 5 percent level, and the J test does not reject the model at any level. This is our preferred regression for the eleven-country panel. Note that the estimates of ρ and γ, 0.0978 and 0.0436, are somewhat larger and smaller, respectively, than those in the literature. Note also that a small value of γ implies high elasticity of intertemporal substitution, $\sigma = 1/\gamma$, and hence high interest sensitivity of household saving.

Using the estimates from this regression, we estimate the derivatives of interest and their approximate standard errors (Kmenta 1971, 444), all

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evaluated at the sample means of the variables. The estimated values of these derivatives have the expected signs (see Part A of table 3), but only the value of $\partial s_{it}/\partial d_{it}$ is statistically significant. Its size, -0.20 (t ratio $= -6.50$), is similar to that found by some studies (King and Dicks-Mireaux 1982; Diamond and Hausman 1984; Hubbard 1986) but is lower than that of other studies (Feldstein and Pellechio 1979; Kotlikoff 1979; Bernheim 1987; Attanasio and Rohwedder 2003).

We now consider again the weak-identification problem discussed earlier, although we have already provided evidence (based on the first-stage R^2 and F statistics) that our GMM regressions do not suffer from the weak-instrument problem. First, we estimate Equation (7) by NLLS. Again, we set $\boldsymbol{x}_{it} = (DG_{it}, IC_{it})'$, as SC_{it} turns out to be insignificant. This is the third regression in Part A of table 2. The estimated coefficients are correctly signed and significant, except for $\hat{\beta}_2$, which is insignificant at conventional levels. Overall, the NLLS estimates differ noticeably from the GMM ones, thus adding credibility to the latter. In particular, the NLLS estimates of ρ, γ, β_1, and β_2 are lower (algebraically) than their GMM counterparts obtained from our preferred regression by 26.9, 31.7, 0.5, and 8.2 percent, respectively.

Second, we drop the two IVs dated $t - 3$ ($s_{it\text{-}3}$ and $gdpe_{it\text{-}3}$) from $\boldsymbol{V}_2$ and reestimate Equation (7) using $M = 29$ IVs ($j/N = 1.7$). The new estimates and their t ratios are as follows (percentage changes in parentheses): $\hat{\rho} = 0.0981$ (0.3), $t_{\hat{\rho}} = 6.36$ (-3), $\hat{\gamma} = 0.0434$ (-0.5), $t_{\hat{\gamma}} = 5.40$ (-7), $\hat{\beta}_1 = -12.60$ (-16.5), $t_{\hat{\beta}_1} = -1.59$ (-39), $\hat{\beta}_2 = 0.83$ (13.7), and $t_{\hat{\beta}_2} =$ 1.44 (-27.3). Note that the estimates of ρ and γ as well as their t ratios are almost the same, those of β_1 and β_2 are somewhat different, and the t-ratios of the latter are noticeably different. The bias/efficiency tradeoff discussed earlier is in operation here, but, fortunately, it concerns only half of the parameters, and the bias, if any, seems to be small.

The Fourteen-country Panel

We follow the same steps here, but use the IV sets $\boldsymbol{V}_1'$ and $\boldsymbol{V}_2'$ ($j/N = 1.7$ in both cases). The results are reported in Part B of tables 2 and 3. $\boldsymbol{V}_2'$ yields our preferred regression, which differs noticeably from that of the eleven-country panel in four ways. First, $\boldsymbol{V}_2'$ contains IVs dated only up to $t - 2$ (so we do not conduct sensitivity analysis here). Second, the value of the coefficient $\hat{\beta}_1$ and its t ratio are much smaller here. Third, the estimated values of the derivatives of interest (table 3) are much larger (in absolute value). Fourth, the values of $\partial p(\boldsymbol{x}_t)/\partial(IC_t)$ and $\partial(\partial s_t/\partial d_t)/\partial(IC_t)$, 0.12 ($t =$

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1.95) and -0.09 ($t = -1.81$), respectively, are now significant at the 5 percent level. The last two findings may reflect the fact that the quality of institutions in this panel is, on average, lower than that in the eleven-country panel, and an improvement has a greater effect. For example, the above two estimates imply that a ceteris paribus increase in the index *IC* by one unit causes (1) the probability of granting pensions to increase by 0.12 and (2) the impact of social security contributions on household saving to become more negative, from -0.83 to -0.92, whereas in the eleven-country panel this effect was insignificant. In other words, the marginal effect of an improvement in the quality of institutions on the credibility of the social security system is greater in countries where the quality of institutions is low than in countries where it is high, a result that stresses the role of institutions in reducing uncertainty about pensions.[12]

The Twenty-five-country Panel

Again, we follow the same steps as before and report the results in Part C of tables 2 and 3. Here, we use the IV sets $\boldsymbol{V}_1''$ ($j/N = 0.76$) and $\boldsymbol{V}_2''$ ($j/N = 0.84$). The latter yields our preferred regression, which differs from the previous ones in three ways. First, the coefficient of *IC*, $\hat{\beta}_2$, is not significant at conventional levels. Second, the estimate of $\partial(\partial s_{it}/\partial d_{it})/\partial(DG_{it})$, 0.73 ($t = 1.33$), is significant at the 10 percent level, implying that a ceteris paribus increase in the debt to GDP ratio by one percentage point causes the impact of social security contributions on household saving to increase from -0.38 to 0.35, a fairly large effect. Third, the NLLS estimates of ρ, γ, β_1, and β_2 differ more noticeably from the GMM ones, specifically by -27.6, -20.2, 37.5, and 21.4 percent, respectively.[13]

Concluding Remarks

In this article, we use a two-period OGM and three panel data sets of annual aggregate data from twenty-five countries to estimate by GMM a fixed-effects Euler equation for household saving. We focus on the effects of some institutional and other variables, such as corruption and the debt to GDP ratio, on the relationship between social security contributions and household saving. These effects are made operational by assuming that these variables determine the probability that a PAYG social security system will grant pensions to the old at retirement.

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We use several tests and robustness checks in an effort to address the well-known problems of weak instruments and weak identification. We find no evidence that the weak-instrument problem is present in our regressions. In addition, with respect to their sign, all of our coefficient estimates are robust to changes in the sample and in the instruments; whereas, with respect to their size, the estimates of the rate of time preference and of the coefficient of relative risk aversion are fairly robust to these changes, but the other coefficient estimates are less robust. The possible biases do not seem to be large, however, so we consider our empirical findings to be reasonably reliable and in accordance with the theoretical implications, hence usable.

First, high levels of corruption and of the debt to GDP ratio reduce the probability that a PAYG system will grant pensions. Second, social security contributions reduce household saving in a less than one-for-one manner. Our estimates of this derivative, evaluated at the sample means of the variables, range from -0.20 to -0.83 (see table 3). Third, there is some evidence that the higher the level of corruption or the debt to GDP ratio, the lower the reduction in household saving caused by an increase in social security contributions, as individuals try to self-insure themselves against the higher uncertainty induced by corruption and indebtedness. The last two findings are important, as they imply a self-insurance behavior, which strengthens saving, thus making the aged less vulnerable to the risk of ending up with lower standards of living at retirement in countries where there is high corruption and indebtedness. Fourth, a ceteris paribus improvement in the quality of institutions has a greater effect on the credibility of the social security system in countries where the quality of institutions is low than in countries where it is high. This result demonstrates the importance of institutional quality in reducing uncertainty surrounding the social security system.

The abovementioned estimates may be useful in evaluating policy proposals aiming to improve the viability of the PAYG social security systems in the countries considered here. Along with the reforms of the PAYG systems that have been taking place in these countries, their governments may be able to improve the viability and credibility of their PAYG systems by reducing corruption and the debt to GDP ratio. They may also find useful to promote the idea of privately funded pensions and, more generally, to encourage private saving, for example, by reducing the tax rates on interest and on capital gains. Such taxation policies are expected to increase the after-tax real interest rate, and hence private saving substantially, since our estimate of the elasticity of intertemporal

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substitution ($1/\gamma$) is quite high (about twenty-five). At times of recession and high external indebtedness in many countries, using appropriate taxation policies that motivate private saving seems crucial in improving national saving and encouraging capital accumulation.

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Authors' Note

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Notes

1. The constant relative risk aversion utility function is required in order for the economy to converge to a balanced growth path (Barro and Sala-i-Martin 1995, 64–65).
2. More generally, an ex-ante real interest rate should be used. Since such a series is not available, however, in the empirical part of the article we are forced to use an ex-post series. This is a consequence of using secondary data (i.e., data collected by others for a different purpose).
3. Equation (4) could be written more generally to include the case of a collapsing social security system (see section 3.2.8 of the second author's doctoral dissertation). Here, we use this special case, because, to our knowledge, in no country of our sample has the system collapsed.
4. The logistic and the standard normal cumulative distribution functions are close to each other, except for the extreme tails. Thus, we are not likely to get different results, unless the samples are large and have enough observations at the tails. Multiplying the estimated coefficients of the probit model by 1.6

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yields approximately the estimated coefficients of the logit model (Maddala 1986, 22–23).

5. Without this simplification, the term u_{t+1} would appear in Equation (6). In the empirical part of this article, it turns out that the presence of the estimated value of u_{t+1} in the derivatives of interest does not affect their sign, so the simplifying assumption $u_{t+1} = 1$ in Equation (5) is empirically justified.
6. In fact, they use a mix of different schemes, but pay-as-you-go is their basic system. These schemes are categorized into two tiers. The first includes public-pension schemes, that is, the basic scheme, which pays flat-rate benefits; the income-tested scheme; and the minimum-pension scheme. The second tier includes the earnings-related and the defined-contribution schemes (OECD 2009).
7. Household saving is the gross saving of households and nonprofit institutions serving households (NPISHs), which are nonmarket producers operating as separate legal entities. Their main resources, apart from property income and occasional sales, are voluntary contributions in cash or in kind made by households and general government. Examples are churches and religious societies, sports and other clubs, and political parties.
8. For Belgium, Canada, Cyprus, Czech Republic, France, Hungary, Italy, Latvia, Lithuania, Poland, Spain, Sweden, the United Kingdom, and the United States, the source of the data for the interest rate is IFS. For Austria, Denmark, Estonia, Finland, Germany, Greece, Japan, the Netherlands, Norway, Portugal, and Slovakia, the source is AMECO.
9. Note that in our data we have $w_{it} - s_{it} - d_{it} > 0$ for every i and t, so that in Equation (7) the variable $c_{it}^* = \ln(w_{it} - s_{it} - d_{it})$ is a finite number for every i and t. Also, we take $e_{t+1} \approx 0$, so that $u_{t+1} = E_t\left[1 \Big/ (1 + e_{t+1})^{1/\gamma}\right] \approx 1$, hence $\ln(u_{t+1}) = u_{t+1}^* \approx 0$, an approximation that is confirmed by the data, that is, in the eleven-, fourteen-, and twenty-five-country panels, the values of the residuals from the preferred regressions range, respectively, from -0.05 to 0.08, from -0.07 to 0.11, and from -0.07 to 0.15.
10. We use zero as a starting value for each parameter. We also tried thirty different combinations of the starting values 0, 0.5, 1, and 1.5 for each parameter but obtained the same estimates.
11. We also used alternative definitions of $\boldsymbol{x}_{it}$, namely, $\boldsymbol{x}_{it} = (DG_{it}, IC_{it}, SC_{it}, PS_{it}, DA_{it})'$, $\boldsymbol{x}_{it} = (DG_{it}, IC_{it}, SC_{it}, PS_{it})'$, $\boldsymbol{x}_{it} = (DG_{it}, IC_{it}, PS_{it}, DA_{it})'$, and $\boldsymbol{x}_{it} = (DG_{it}, IC_{it}, SC_{it}, DA_{it})'$, and alternative instrumental variable (IV) sets. These alternatives failed to yield statistically significant and correctly signed coefficients, however.

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12. In a previous version of this article, we did not estimate Equation (7) for the fourteen-country panel. An anonymous referee of this journal suggested that we do this estimation and make such comparisons.
13. As in the fourteen-country panel, the "deepest" lag contained in the IV set V_2'' is $t - 2$, so, again, we do not conduct sensitivity analysis (defined at the end of the fourth section). Note, however, that if we drop the eight IVs dated $t - 2$, and estimate Equation (7) with $M = 37$ IVs ($j/N = 0.52$), the estimates of ρ, γ, β_1, and β_2 are, respectively, 0.0901 ($t = 4.84$), 0.0573 ($t = 3.58$), -0.72 ($t = -0.59$), and 0.35 ($t = 0.29$).

Supplemental Material

The online appendices are available at http://pfr.sagepub.com/supplemental.

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