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Valuing Lost Home Production for Dual-Earner Couples

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Valuing Lost Home Production in Dual-Earner Couples

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Abstract

Economists' principal tool for studying household behavioral responses to changes in tax and other government policies, and the magnitude and determinants of private saving, is the life—cycle model. The purpose of this paper is to attempt to incorporate into that model one of the most conspicuous changes in the U.S. economy in the last 50 years, the rise in labor market participation for married women. The increased presence of married women in the labor force has obvious benefits: women now earn much more income than they did in the past. On the other hand, working women presumably spend less time doing housework and other types of home production, and the forgone value of time at home reduces the net benefit of their work in the market. Conventional accounts do not provide measurements of the costs of lost home production, but we attempt to use comparisons of household net worth at retirement to deduce valuations indirectly. This paper modifies a standard life—cycle model to include women's labor supply decisions, estimates key parameters of the new specification, and attempts to assess the significance of rising female labor market participation for aggregate national saving in the U.S. Using panel data from the Health and Retirement Study, we find that the difference between measured labor market earnings for married women and earnings net of the value of lost home production seems moderately small – about 30 percent – and that the corresponding long—run effect on the overall rate of private saving is minor.

Valuing Lost Home Production for Dual–Earner Households

John Laitner, Dmitriy Stolyarov, Chris House

Economists' principal tool for studying households' behavior, including likely household responses to changes in tax and other government policies, and including the magnitude and determinants of private saving, is the life–cycle model of Diamond [1965], Auerbach and Kotlikoff [1987], and, in particular, Modigliani [1986].[1] The purpose of this paper is to attempt to incorporate into that model one of the most conspicuous and important changes in the U.S. economy in the last 50 years, the rise in labor market participation for married women.[2] Goldin [1990, p.10], for example, shows that the participation rate in the labor force for all women grew from 19 percent in 1890 to 60 percent in 1990 and that the participation rate for married women rose even more, from 5 percent in 1890 to nearly 60 percent in 1990. By 2002, in fact, the labor force participation rate for married women was 67 percent. This paper modifies a standard life–cycle model to include women's labor supply decisions, estimates key parameters of the new specification, and attempts to assess the significance of rising female labor market participation for aggregate national saving in the U.S.

In general, technological progress tends to push the standard of living for U.S. households steadily upward, and in this context it is somewhat surprising to find a trend away from home production — in other words, to find households, despite their growing prosperity, willingly giving up the amenities of home production of services such as food preparation, housekeeping, and childcare. If we are to bring female labor force participation into the life–cycle model, we must be able to analyze the tradeoff that households face: as a woman joins the labor force, she raises her household's earnings; however, as she devotes more hours to market work, she almost surely reduces her time allocation to home production. This paper first presents a life–cycle model that incorporates the time allocation decision between home production and market work that married women, and their families, face.

It is difficult to measure the value of home production: survey data and macroeconomic time series (including the National Income and Product Accounts) rely on recording market transactions; so, conventionally they omit home production. We have, consequently, directly available only part of the information that we would like for the study household choices: household earnings (and GDP) fully reflect increases in female labor force participation, but there is little data on corresponding diminutions in home production. This paper follows, therefore, an indirect approach: it attempts to use its theoretical framework in combination with microeconomic data from the Health and Retirement Study (HRS) to estimate *net* gains in household resources (i.e., additional market earnings less the value of sacrificed home production) when women join the labor force — distinguishing net

[1] Recent examples include, for instance, Hubbard *et al.* [1994], Altig *et al.* [2001], Gokhale *et al.* [2001], Laitner [2001].

[2] For other recent work, see Benhabib *et al.* [1991], Greenwood *et al.* [1995], and Rupert *et al.* [2000].

from gross resource gains for individual households on the basis of size of the households' subsequent accumulations of wealth for their retirement.

Finally, employing the new model and parameter estimates, we simulate the effect of increased female labor market participation on the average national savings rate. A feature of this paper is that savings behavior enters our analysis twice: first, our strategy for identifying the value of lost home production for women who join the labor force relies upon microeconomic data on savings behavior; second, we simulate the macroeconomic effect of women's labor market participation on the U.S. aggregative average propensity to save.

The following examples illustrate our strategy for obtaining measurements of the value of changes in home production, and they also suggest why we think that female labor force participation could influence a nation's overall saving rate. (i) In a "traditional" household, suppose a husband works two–thirds of his adult life, earning \$900,000, and his wife never does market work. Suppose the household saves \$300,000 for retirement — seeking to hold the level of its consumption constant. (ii) In a second traditional household, the husband alone works outside of the house — for two–thirds of his adult life — and he earns \$1,800,000. Assuming preference orderings are homothetic, suppose the second household saves \$600,000 for its retirement. (iii) In a third, "modern" household, the husband works for two–thirds of his adult life and earns \$900,000, and his wife works outside of the house and earns \$900,000. As she participates in the labor market, the wife drastically cuts back on her home production. The household purchases market replacements for the lost services. Suppose the replacements cost \$900,000. Since the household's "net" earnings are the same as the first household's, suppose that it saves \$300,000 for retirement, too. This paper's analysis depends on a comparison of such varying cases. In particular, we argue that the fact that accumulated saving in the last case may resemble that of household (i) rather than (ii) can provide information on the magnitude of lost home production. Our reasoning is that households with the same net–of–lost–home–production earnings should want the same net worth at retirement; thus, similar observed net worth for two households and dissimilar gross earnings implies the need for home–production valuations that bring net earnings into equality. Conversely, from a macroeconomic perspective, the fact that case (iii) has recently become more widespread suggests one reason why the ratio of national savings to earnings may have fallen: a standard measure of national saving is the ratio of aggregative saving to aggregative gross output; however, life–cycle theory tends to imply that the ratio of aggregative savings to "net" earnings should remain constant — and the discrepancy between aggregative net (of sacrificed home production) and gross earnings has presumably grown in recent decades.

The organization of this paper is as follows. Sections 1-2 present our life–cycle model. Section 3 discusses our data. Section 4 presents our parameter estimates. Section 5 considers the possible impact of increased female labor force participation on the aggregative average propensity to save. Section 6 concludes.

1. The Model

Sections 1-2 present a life–cycle model of an individual household's behavior. We show that the model leads to a regression equation

$$\ln(NW_{is}) = \ln\big(\kappa(s,\sigma)\big) + \ln\big(Y^M_{is} + (1-\theta)\cdot Y^F_{is}\big) + \epsilon_{is}\,, \tag{1}$$

where NW_{is}, Y^M_{is}, and Y^F_{is} are, respectively, net worth for household i at age s, the present value at household age s of the lifetime earnings of the household's adult male, and the present value of adult female's lifetime earnings; where $\theta \in (0,1)$ and σ are parameters that we will estimate; where $\kappa(.)$ is a known function; and where ϵ_{is} is a regression error reflecting, say, measurement error in NW_{is}.

Specification

Focus on a single household that lives over ages $t = 0$ to $t = T$. It has market consumption c_t at age t. The household includes a man and wife. The man earns y^m_{it}, after taxes, at age t, inelastically supplying labor. Until retirement, the wife works full-time as well. However, she divides her work time between labor force participation and home production. Let $h^f_t \in [0,1]$ be the fraction of the week that the woman works in the labor market, and let $1 - h^f_t$ be the fraction of her work-time that she devotes to home production. She earns an aftertax wage rate w^f_{it} for each week of market work. At exogenously specified age R (but see below), the household retires, y^m_{it} drops to zero, and so does h^f_t. The aftertax market real interest rate is r.

Household i determines its life–cycle behavior from

$$\max_{c_t \geq 0,\, h^f_t \geq 0} \int_0^T e^{-\rho\cdot t}\cdot u\big(c_t - A_{it}\cdot [h^f_t]^{\xi}\big)\,dt\,, \tag{2}$$

$$\text{subject to: } \dot{a}_t = r\cdot a_t + y^m_{it} + h^f_t\cdot w^f_{it} - c_t\,,$$

$$a_0 = 0 \quad \text{and} \quad a_T \geq 0\,,$$

$$h^f_t = 0 \quad \text{and} \quad y^m_{it} = 0 \quad \text{all} \quad t \geq R\,.$$

The household's net worth at age t is a_t.

Following much of the literature, assume isoelastic preferences:

$$u(x) \equiv \frac{[x]^{\gamma}}{\gamma}\,, \quad \gamma < 1\,. \tag{3}$$

Assume $A_{it} > 0$ is exogenously given and

$$\xi > 1\,. \tag{4}$$

In the terminology of equation (1),

$$NW_{is} = a_s\,, \quad Y^M_{is} = \int_0^R e^{-r\cdot(t-s)}\cdot y^m_{it}\,dt\,, \quad Y^F_{is} = \int_0^R e^{-r\cdot(t-s)}\cdot h^f_t\cdot w^f_{it}\,dt\,. \tag{5}$$

Section 2 discusses the precise link from model (2) to regression equation (1).

Discussion of the Specification

We close this section with observations about the generality of our specification (though we leave until the end of Section 2 a discussion of the asymmetry of our treatment of men and women).

Heterogeneity. Our specification allows a great deal of heterogeneity among households — i.e., men can have different earning abilities, y^m_{it}; women can have different market opportunities w^f_{it}; and women can have different efficiencies in home production, A_{it}. The model, therefore, predicts that some married women will work in the labor market more hours than others. For example, in comparing two households, if the man in the second earns twice as much, his wife has twice the wage rate, and his wife is twice as good at home production, the model predicts that women in both households will supply the same number of market hours. In contrast, if the man in the second household earns twice as much, his wife has twice the wage rate, but his wife is no better than the first wife at home production, the model predicts that the woman in the second household will allocate more hours to labor market participation. We could elaborate our utility function to incorporate differing numbers of children at different ages (e.g., Laitner and Silverman [2005]), and A_{it} could be higher when more children are at home. The role of children remains a topic for future work.

Endogenous Retirement. Although we assume that retirement age is exogenously given, future work will endogenize it (see, for instance, Laitner and Silverman [2005]). In the meantime, our regression analysis below does take the possible endogeneity of R_i into account.

Time–Allocation Limits. One might want to impose a constraint

$$h^f_{it} \in [\breve{h}\,,\,\bar{h}] \subset (0,1)\,. \tag{6}$$

The idea of a lower limit $\breve{h} > 0$ might be that there are fixed costs to having a market job at all — for both an employee and an employer — so that labor force participation with very short hours almost never occurs. This paper ignores such a possibility — in part because very low female earnings have little weight in determining θ in equation (1).

The idea of an upper bound might be that employers often must, by statute, pay overtime wages for a workweek exceeding 40 hours and are reluctant to do so. In practice, however, a worker might take a second job. What seems at least as likely is that an ambitious employee might work extra hours per week "off the clock" in order to secure future salary raises, promotions, etc.[3]

Perhaps a more plausible obstacle constraining time allocation is as follows: due to coordination problems with other workers, efficiency costs to frequent startups, and advantages to finishing tasks promptly without requiring many workers to incur learning costs for the same project, employers might in practice offer a higher wage rate to full-time employees than to part-time workers. In other words, for some $\bar{h}$, we might have

[3] A dynamic model of human capital, including experience, accumulation is beyond the scope of this paper and remains a topic for future research. See, for example, Ryder *et al.* [1976] and Attanasio *et al.* [2004].

$$w_{it}^f = \begin{cases} w_{it}^0 \, , & \text{if } h_{it}^f \geq \bar{h} \\ w_{it}^{00} < w_{it}^0 \, , & \text{if } h_{it}^f < \bar{h} \end{cases} . \tag{7}$$

This could, for instance, lead a woman who would otherwise prefer a market job for 25 hours per week to select a 40-hour per week job instead (because of a lower wage rate for the former), and it would complicate our solution in the next section. Our data set (see below) does not provide information on work hours (before 1992), and we leave this issue as a topic for future research.

2. Solution of the Model

Model (2) presents a concave dynamic maximization problem with convex constraints; therefore, the conventional first–order conditions of optimal control theory determine the unique solution. Drop the subscript i for convenience in this section.

Define

$$z_t \equiv c_t - A_t \cdot [h_t^f]^{\xi} \, . \tag{8}$$

Then one can transform model (2) into

$$\max_{z_t \geq 0, \, h_t^f \geq 0} \int_0^T e^{-\rho \cdot t} \cdot u(z_t) \, dt \, , \tag{9}$$

$$\text{subject to: } \dot{a}_t = r \cdot a_t + y_t^m + h_t^f \cdot w_{it}^f - A_t \cdot [h_t^f]^{\xi} - z_t \, ,$$

$$a_0 = 0 \quad \text{and} \quad a_T \geq 0 \, ,$$

$$y_t^m \quad \text{and} \quad h_t^f = 0 \quad \text{all} \quad t \geq R \, .$$

The transformed problem has current-value Hamiltonian

$$\mathcal{H} = u(z_t) + \lambda_t \cdot \big[r \cdot a_t + y_t^m + h_t^f \cdot w_t^f - A_t \cdot [h_t^f]^{\xi} - z_t \big] \, . \tag{10}$$

The following conditions are necessary and sufficient for an optimum:

$$\frac{\partial \mathcal{H}}{\partial z_t} = 0 \Longleftrightarrow u'(z_t) = \lambda_t \tag{i}$$

$$\frac{\partial \mathcal{H}}{\partial h_t^f} = 0 \Longleftrightarrow \lambda_t \cdot \big[w_t^f - \xi \cdot A_t \cdot [h_t^f]^{\xi - 1} \big] = 0 \tag{ii}$$

$$\dot{\lambda}_t = \rho \cdot \lambda_t - \frac{\partial \mathcal{H}}{\partial a_t} \Longleftrightarrow \dot{\lambda}_t = \lambda_t \cdot [\rho - r] \, , \tag{iii}$$

$$\dot{a}_t = r \cdot a_t + y_t^m + h_t^f \cdot w_t^f - A \cdot [h_t^f]^{\xi} - z_t \, , \tag{iv}$$

$$\lambda_T \geq 0 \quad \text{and} \quad \lambda_T \cdot a_T = 0\,, \tag{v}$$

$$a_0 = 0\,. \tag{vi}$$

Consider the optimum. Conditions (i) and (iii) show $\lambda_t > 0$ all t; thus, (v) implies

$$a_T = 0\,. \tag{11}$$

From (i),

$$\frac{[z_t]^{\gamma-1}}{\gamma} = \lambda_t\,.$$

Taking logs and then time derivatives,

$$(\gamma - 1) \cdot \dot{z}_t / z_t = \dot{\lambda}_t / \lambda_t\,.$$

Using (iii),

$$\dot{z}_t / z_t = (r - \rho)/(1 - \gamma)\,.$$

Suppose that the household is retired at age s. Integrating both sides from 0 to s,

$$z_s = z_0 \cdot e^{\frac{r-\rho}{1-\gamma} \cdot s}\,. \tag{12}$$

Multiply every term in (iv) by $e^{-r\cdot(t-s)}$. Then

$$\begin{aligned} \frac{d(e^{-r\cdot(t-s)} \cdot a_t)}{dt} &= e^{-r\cdot(t-s)} \cdot [\dot{a}_t - r \cdot a_t] \\ &= e^{-r\cdot(t-s)} \cdot \left[y_t^m + h_t^f \cdot w_t^f - A_t \cdot [h_t^f]^\xi - z_t\right]. \end{aligned} \tag{13}$$

Integrating the first and last expressions from $t = s \geq R$ to $t = T$, using the fundamental theorem of calculus, noting equation (11), and noting that $R \leq s$ means the first three right–hand side terms are zero, we have

$$-a_s = -\int_s^T e^{-r\cdot(t-s)} \cdot z_t \, ds \Longleftrightarrow a_s = \int_s^T e^{-r\cdot(t-s)} \cdot z_t \, dt\,. \tag{14}$$

On the other hand, integrating the extreme left and right sides of (13) from $t = 0$ to $t = T$, we have

$$\begin{aligned} 0 - 0 &= \int_0^T e^{-r\cdot(t-s)} \cdot [y_t^m + h_t^f \cdot w_t^f - A_t \cdot [h_t^f]^\xi - z_t] \, dt \\ \Longleftrightarrow Y_s^M + Y_s^F - &\int_0^T e^{-r\cdot(t-s)} \cdot A_t \cdot [h_t^f]^\xi \, dt = \int_0^T e^{-r\cdot(t-s)} \cdot z_t \, dt\,. \end{aligned} \tag{15}$$

Since $\lambda_t > 0$ all t, condition (ii) implies

$$h_t^f = [\frac{w_t^f}{\xi \cdot A_t}]^{\frac{1}{\xi - 1}} . \tag{16}$$

Then

$$A_t \cdot [h_t^f]^{\xi} = A_t \cdot [\frac{w_t^f}{\xi \cdot A_t}]^{\frac{\xi}{\xi-1}} = [A_t]^{\frac{-1}{\xi-1}} \cdot [w_t^f]^{\frac{\xi}{\xi-1}} \cdot [\xi]^{\frac{-\xi}{\xi-1}} ,$$

and

$$h_t^f \cdot w_t^f = [\frac{w_t^f}{\xi \cdot A_t}]^{\frac{1}{\xi-1}} \cdot w_t^f = [w_t^f]^{\frac{\xi}{\xi-1}} \cdot [\xi]^{\frac{-1}{\xi-1}} \cdot [A_t]^{\frac{-1}{\xi-1}} ;$$

hence,

$$A_t \cdot [h_t^f]^{\xi} = \theta \cdot h_t^f \cdot w_t^f , \tag{17}$$

where

$$\theta \equiv [A_t]^{\frac{-1}{\xi-1}} \cdot [\xi]^{\frac{-\xi}{\xi-1}} \cdot [\xi]^{\frac{1}{\xi-1}} \cdot [A_t]^{\frac{1}{\xi-1}} = \frac{1}{\xi} . \tag{18}$$

Note that since $\xi > 1$, we have

$$\theta \in (0, 1) . \tag{19}$$

Returning to (15), substitution from (17) yields

$$\begin{aligned} Y_s^M + Y_s^F - \int_0^T e^{-r \cdot (t-s)} \cdot A_t \cdot [h_t^f]^{\xi} \, dt \\ = Y_s^M + Y_s^F - \theta \cdot Y_s^F = \int_0^T e^{-r \cdot (t-s)} \cdot z_t \, dt . \end{aligned} \tag{20}$$

We are now ready to deduce equation (1). Continue to assume that the household that we are considering is retired — i.e., $s \geq R$. Using (5), (14), and (20), we have

$$\frac{NW_s}{Y_s^M + (1 - \theta) \cdot Y_s^F} = \frac{\int_s^T e^{-r \cdot (t-s)} \cdot z_t \, dt}{\int_0^T e^{-r \cdot (t-s)} \cdot z_t \, dt} .$$

After substituting from (12), cancel $e^{r \cdot s}$ and z_0 from the top and bottom on the right side. Define

$$\sigma \equiv -r + \frac{r - \rho}{1 - \gamma} . \tag{21}$$

Then

$$\frac{NW_s}{Y_s^M + (1-\theta) \cdot Y_s^F} = \frac{\int_s^T \; e^{\sigma \cdot t} \, dt}{\int_0^T \; e^{\sigma \cdot t} \, dt} \, . \tag{22}$$

Calling the ratio on the right–hand side $\kappa(s, \sigma)$, we have

$$\frac{NW_s}{Y_s^M + (1-\theta) \cdot Y_s^F} = \kappa(s, \sigma) \, . \tag{23}$$

Taking logs of both sides, we have equation (1).

Discussion of Result

Given homotheticity of preferences, a natural outcome to expect might be that for a given s,

$$\frac{NW_s}{Y_s^M + Y_s^F} = \text{constant} \, . \tag{24}$$

For a household with a non-working wife, (23) and (24) are identical. With a working wife, the two are different. Indeed, our model implies that the left–hand side of (24) is not constant — rather it will decline with increases in Y_s^F. The National Income and Product Accounts omit home production from output. Once one neglects the value of home production, it is easy to overlook the consequences of its absence. Our analysis implies that such a mind set will tend to cause one to perceive a drop in private wealth accumulation during periods in which women reduce hours of work at home in preference for market jobs: if $Y_s^F = 0$, we have

$$\frac{NW_s}{Y_s^M + Y_s^F} = \frac{NW_s}{Y_s^M + (1-\theta) \cdot Y_s^F} = \kappa(s, \sigma) \, ;$$

however, when $Y_s^F > 0$, we predict

$$\frac{NW_s}{Y_s^M + Y_s^F} < \frac{NW_s}{Y_s^M + (1-\theta) \cdot Y_s^F} = \kappa(s, \sigma) \, .$$

Men and Women's Roles

Our analysis assumes a set workweek for men and for women (perhaps emerging from a labor/leisure choice beyond the scope of this paper); it investigates the allocation of time within that workweek between labor force participation and home production. In the mathematics above, the time allocation problem — for women — is static: essentially, at each household age t, a woman chooses h_t^f to satisfy

$$\max_{h_t^f \in [0,1]} \{ w_t^f \cdot h_t^f - A_t \cdot [h_t^f]^{\xi} \} \, . \tag{25}$$

To treat men and women more symmetrically, we could make household utility

$$u\big(c_t - [A_t^f \cdot h_t^f + A_t^m \cdot h_t^m]^\xi\big) \tag{26}$$

and male earnings $w_t^m \cdot h_t^m$, with $h_t^m \in [0,1]$ the fraction of a male's workweek allocated to home production. If childhood training, education, and cultural biases imply $A_t^m \leq A_t^f$ and $w_t^m \geq w_t^f$, we could argue that the outcome will always be $h_t^m = 1$. This provides a possible justification for our asymmetric treatment of men and women above.

A justification that we prefer, however, is as follows. Suppose that men and women are not perfect substitutes in home production. In fact, let household utility be

$$u\big(c_t - A_t^f \cdot [h_t^f]^{\xi^f} - A_t^m \cdot [h_t^m]^{\xi^m}\big)\,, \quad \xi^f > 1\,, \quad \xi^m > 1\,. \tag{27}$$

Suppose that in practice, both sexes devote time to both labor force participation and home production. Solving the analogue of (25) for men,

$$\frac{\text{home production}}{\text{market earnings}} = \frac{A_t^m \cdot [h_t^m]^{\xi^m}}{h_t^m \cdot w_t^m} = \frac{A_t^m \cdot [h_t^m]^{\xi^m}}{h_t^m \cdot \xi^m \cdot A_t^m \cdot [h_t^m]^{\xi^m - 1}} = \frac{1}{\xi^m}\,. \tag{28}$$

Similarly, for women,

$$\frac{\text{home production}}{\text{market earnings}} = \frac{A_t^f \cdot [h_t^f]^{\xi^f}}{h_t^f \cdot w_t^f} = \frac{A_t^f \cdot [h_t^f]^{\xi^f}}{h_t^f \cdot \xi^f \cdot A_t^f \cdot [h_t^f]^{\xi^f - 1}} = \frac{1}{\xi^f}\,. \tag{29}$$

Our data below covers households retiring 1990-2005. For that cohort, casual observation strongly suggests a much higher ratio of value of home production to market earnings for women than for men. On the basis of this observation, we might assume

$$\frac{1}{\xi^f} > \frac{1}{\xi^m}\,. \tag{30}$$

A symmetric treatment of men and women stemming from (27) would lead to a version of (1) with

$$\begin{aligned} \ln(NW_s) &= \ln\big(\kappa(s,\sigma)\big) + \ln\big((1 - \frac{1}{\xi^m}) \cdot Y_s^M + (1 - \frac{1}{\xi^f}) \cdot Y_s^F\big) + \epsilon_t \\ &= \ln\big(\bar{\kappa}(s,\sigma)\big) + \ln\big(Y_s^M + (1 - \bar{\theta}) \cdot Y_s^F\big) + \epsilon_s \end{aligned} \tag{31}$$

where

$$1 - \bar{\theta} \equiv \frac{1 - \frac{1}{\xi^f}}{1 - \frac{1}{\xi^m}}\,,$$

$$\begin{aligned} \ln\big(\bar{\kappa}(s,\sigma)\big) &\equiv \ln\big(\kappa(s,\sigma)\big) + \ln(1 - \frac{1}{\xi^m}) \\ &= \ln\big(\kappa(s,\sigma) \cdot (1 - \frac{1}{\xi^m})\big)\,. \end{aligned}$$

Since

$$\frac{1}{\xi^m} < \frac{1}{\xi^f} \Longleftrightarrow 1 - \frac{1}{\xi^m} > 1 - \frac{1}{\xi^f} \Longleftrightarrow 1 > \frac{1 - \frac{1}{\xi^f}}{1 - \frac{1}{\xi^m}}$$

and the last expression is positive, we have $\bar{\theta} \in (0, 1)$. As we estimate θ from equation (1), one could interpret our value as actually determining $\bar{\theta}$ from (31).

In addition to treating males and females symmetrically, the second justification has several advantages. First, it recognizes the fact that men, as well as women, provide some home production. Second, it is compatible with a comprehensive definition of the workweek. For example, a woman participating in the labor market 40 hours per week could be providing 20 hours per week of home production as well, for a total workweek of 60 hours. When our model can encompass such a long workweek, even "full-time" participation in the labor market is fully compatible with our first–order condition (e.g., Hamiltonian first–order condition (ii)) for home production.

3. Data

We use data from the Health and Retirement Study (HRS) to predict lifetime earnings of men and women, and household net worth. We assume a gross–of–tax real interest rate of 5 percent per year.[4] Our income tax rate, $\tau = .15$, comes from government spending on goods and services less indirect taxes.[5] Thus, we set $r = .05 \cdot (1.0 - .15) = 0.0425$. This paper's focus is married couples.

Household Net Worth

We use the original survey cohort from the HRS, consisting of households in which the respondent is age 51-61 in 1992.[6] In each survey wave (i.e., 1992, 1994, 1996, 1998,

[4] Suppose that we calibrate our real interest rate from a ratio of factor payments to capital over the market value of private net worth. For the numerator, NIPA Table 1.13 gives corporate business income, indirect taxes, and total labor compensation. The first less the other two is a measure of corporate profits (net of depreciation); the ratio of profits to profits plus labor remuneration is "profits share." Multiply the latter times corporate and noncorporate business income plus nonprofit–institution income, less indirect taxes. Add the income of the household sector (see NIPA Table 1.13) less indirect taxes and labor remuneration. Finally, reduce the numerator by personal business expenses (brokerage fees, etc. from NIPA Table 2.5.5, rows 61–64). For the denominator, use U.S. Flow of Funds household and private non–profit institution net worth (Table B.100, row 19), less government liabilities (Table L106c, row 20). Average the net sum at the beginning and end of each given year. The average ratio 1952–2003 of the numerator to the denominator is .0504.

[5] Dividing by national income, the average over 1952–2003 is 14.28%/year.

[6] Age in this case refers to the age of the household's "principal respondent." For a couple, the "principal respondent" can be either male or female — and the survey incorporates the principal respondent's spouse.

2000, 2002), the HRS obtains a complete inventory of each household's assets (including own home, other real estate, automobiles, bank accounts, stocks and bonds, equity in own business, and equity in insurance) as well as debts. Prior to retirement assets include value of defined contribution pension accounts but not the capitalized value of future defined benefit pension rights; thus, we restrict our analysis to the net worth of retired couples. After retirement, the HRS asks households about their pension and Social Security Benefit flows — collecting information separately for husbands and wives on up to three pensions, three annuities, and Social Security Benefits.[7]

For our regress in Section 4, we exclude couples with either member under age 50, with males over 74, and with females over 80.[8] The median male retirement age 1990-2000 is 62, and we restrict our sample to males who retire at ages 56-68 and to couples with 6 years or less difference in age (see below).

Table 1. Distribution of Household Net Worth: 1984 Dollars (NIPA PCE Deflator); HRS Household Weights

Variable	All FINR:[a]	Retired Couples:[b]		
	HRS Net Worth	HRS Net Worth	Add Private Pensions	Add SSB
Minimum	-4,494,000	-41,000	-31,000	39,000
Lower Quart.	21,000	91,000	149,000	252,000
Median	76,000	198,000	281,000	398,000
Upper Quart.	194,000	392,000	488,000	588,000
Maximum	5,842,000	4,615,000	4,808,000	4,935,000
Mean	193,000	314,000	389,000	494,000
Coef. Var.	3.1	1.3	1.1	0.9
Observations	40024	795	795	795
No. Households	7612	388	388	388

a. Sample is households of all "financial respondents" with valid net worth from original sample HRS.
b. Sample is above intersection married couple with both spouses alive, husband at least one earning observation, currently retired, retirement age 56-68, 9-24 years of education, current age 50-74; wife never worked or retired, with linked Social Security earning record, 9-24 years education, current age 50-80. See text.

[7] For the first two pensions, and the first two annuities, the HRS collects data on whether the flow is real or nominal, and on whether the flow carries survivorship rights. Our calculations assume a nominal interest rate three percent higher than our real rate.

[8] As explained below, our model assumes males die when they turn 75, and females when they turn 81. Very young retirees are presumably independently wealthy or disabled — both outside the scope of our model.

Table 1 presents details on our sample's net worth. Clearly private pensions and Social Security Benefits are important components of net worth at ages near retirement.

Male Lifetime Earnings

The HRS links to their Social Security Administration (SSA) earnings histories participants who sign permission waivers. An individual's "earnings history" includes annual data on his/her number of quarters with Social Security coverage and his/her Social Security taxable earning amount. Each history runs 1951–91. The HRS survey waves for 1992, 1994, 1996, 1998, 2000, and 2002 collect previous year's earnings (but see below) and hours for all individuals. Two great advantages of the SSA linked data are that (i) even for a 61 year old in 1992, the SSA history includes earnings back to age 20, and (ii) the SSA data provides administrative-record quality observations. Disadvantages are (a) the Social Security System does not cover all jobs; (b) Social Security records do not include work hours;[9] and, (c) histories are right–censured in some cases because they only track earnings up the to year's statutory maximum subject to the Social Security tax.[10]

Our procedure for characterizing male lifetime earnings is as follows. We are willing to assume that men typically work full time. Our SSA linked earnings histories do not cover all jobs (especially in early years) and are subject to right–censuring in a large number of cases; thus, from all available positive (see below) male earnings observations, we estimate an earnings dynamics model, using a random–effects panel data specification and maximum likelihood methodology that takes account of censuring, and we use the model to impute male earnings prior to 1992. After 1992 (when part–time work prior to retirement might occur), we employ earnings from the survey. Appendix I provides a detailed explanation.

Since our data registers take home pay, we multiply earnings at each age by the year's ratio of NIPA total compensation to NIPA wage and salary accruals. Then we subtract employee and employer Social Security taxes and income taxes at our constant rate of .15.

Table 2 presents information on the distribution of male lifetime earnings. Column 1 includes all men who were married at one time, had at least one annual earnings figure that passed our screens (see Appendix I), had 9-24 years of education and a valid retirement age (see Appendix I), and had a retirement age between the ages of 50 and 69. Column 2 further restricts the sample to men who have at least one earnings observation and appear in Table–1 households.

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[9] Although the histories do annually report quarters of coverage, the definition of a "covered quarter" changes over time and is not directly related to work hours (e.g., Social Security Administration [2002, tab 2.A7]).

[10] The SSA also provides linked W2 tax reports annually 1980-91. Although the W2 records are right–censured for confidentiality, the upper limit is substantially higher than the Social Security earnings cap — $125,000 for earnings under $250,000, $250,000 for earnings under $500,000, and $500,000 for earnings above that amount. In practice, we assume right–censoring at $125,000 for all W2 amounts at or above $125,000. The W2 amounts include non–FICA earnings — and separately identify the latter. They omit some tax deferred pension amounts. Although they also omit self–employment earnings, they identify Social Security measures of the latter. In practice, an individual may have multiple jobs, and we add the corresponding W2 amounts.

Table 2. Distribution of Present Value at Age 50 Lifetime Male Earnings: Gross of Benefits; Net of Taxes; 1984 Dollars (NIPA PCE Deflator); HRS Household Weights		
Variable	Large Sample[a]	Restricted Sample[b]
Minimum	199,000	448,000
Lower Quartile	965,000	1,062,000
Median	1,309,000	1,403,000
Upper Quartile	1,750,000	1,802,000
Maximum	11,181,000	7,432,000
Mean	1,482,000	1,557,000
Coefficient of Variation	1.5	0.5
Observations	2692	388
Number of Households	2692	388

a. Sample is males with 9-24 years of education, age 50-74, at least one earn observation.
b. Sample is one male per household from column 2, Table 1.

Female Lifetime Earnings

We are unwilling to assume that women necessarily work in the labor force full time until retirement. As for men, we estimate a conventional earnings dynamics equation and use it to impute replacements for right–censored observations. We take all other observations, including zeros, directly from the data. Appendix I provides details.

Since we are not imputing substitutes for zeros (or very low reported earnings figures), we make an additional correction to the data as follows. We re–estimate our earnings dynamics equation using all positive earnings values (even the lowest). If a respondent reported that she had previous years of non–FICA employment and the number of years, we impute such earnings using the second earnings dynamics equation. See Appendix I for details.

Finally, as in the case of males, we multiply earnings at each age by the year's ratio of NIPA total compensation to NIPA wage and salary accruals. Then we subtract employee and employer Social Security taxes and income taxes at our constant rate of .15.

Table 3, column 1, presents information on the distribution of lifetime earnings for women with 9-24 years of education, a usable retirement age, linked SSA data (or a survey response indicating no labor force participation prior to 1992), and a retirement age 50-69. As with men, all figures are present values at respondent age 50 and 1984 dollars. Column 2 restricts the sample as in Table 1. Column 3 recomputes column–2 present values at age 50 for the husband (see below).

Table 3. Distribution of Present Value Female Lifetime Earnings: Gross of Benefits; Net of Tax; 1984 Dollars (NIPA PCE Deflator); HRS Household Weights			
Variable	Large Sample[a]	Restricted Sample[b]	
		PV Wife Age 50	PV Husband Age 50[c]
Minimum	0.0	0.0	0.0
Lower Quartile	133,000	186,000	165,000
Median	329,000	338,000	310,000
Upper Quartile	610,000	584,000	510,000
Maximum	2,969,000	1,386,000	1,365,000
Mean	421,000	410,000	338,000
Coef. Variation	0.9	0.7	0.7
Observations	2582	388	795
Number of Households	2582	388	795

a. Sample is females with 9-24 years of education, age 50-74, never worked or linked Social Security earning records, valid retirement age.
b. Sample is one female per household from column 2, Table 1.
c. Sample is one female per wave from column 2, Table 1. (Recall that age varies by wave.)

4. Regression Outcomes

We estimate $\theta = 1/\xi$ from equation (22) using our HRS data, including linked Social Security earnings records. We use a subsample consisting of married couples with both spouses surviving yet retired. The latter enables us to compute the capitalized values of household private pensions, as described in Section 3. We assume that all households have the same parameters ξ, γ, and ρ; hence, θ and σ are the same for all families.

For household i of age s, with $s \geq R_i$, equation (22) is

$$\frac{NW_{is}}{Y_{is}^M + (1-\theta) \cdot Y_{is}^F} = \frac{\int_s^T \, e^{\sigma \cdot t} \, dt}{\int_0^T \, e^{\sigma \cdot t} \, dt} \, . \tag{22'}$$

Note that for the denominator on the left–hand side we need both Y^M and Y^F in present value at the same date — i.e., household age s. Henceforth, we take a household's age to be the husband's age. It is convenient to have all lifetime earnings at a common household age, namely, 50; so, we use the fact that

$$\frac{NW_{is} \cdot e^{-r \cdot (s-50)}}{Y_{i,50}^M + (1-\theta) \cdot Y_{i,50}^F} = \frac{NW_{is}}{Y_{is}^M + (1-\theta) \cdot Y_{is}^F} \, .$$

Similarly, we rewrite the right–hand side of (22') as follows:

$$\frac{\int_s^T e^{\sigma\cdot t}\,dt}{\int_0^T e^{\sigma\cdot t}\,dt} = \frac{e^{\sigma\cdot s}\cdot\int_s^T e^{\sigma\cdot(t-s)}\,dt}{\int_0^T e^{\sigma\cdot t}\,dt} = \frac{e^{\sigma\cdot(s-50)}\cdot\int_s^T e^{\sigma\cdot(t-s)}\,dt}{e^{-\sigma\cdot 50}\cdot\int_0^T e^{\sigma\cdot t}\,dt}$$
$$= \frac{e^{\sigma\cdot(s-50)}\cdot\int_0^{T-s} e^{\sigma\cdot v}\,dv}{e^{-\sigma\cdot 50}\cdot\int_0^T e^{\sigma\cdot t}\,dt}\,. \tag{32}$$

Since the denominator of this term is independent of i and s, we fold it into the regression constant. It is easy to perform the integration in the numerator — i.e.,

$$\int_0^{T-s} e^{\sigma\cdot v}\,dv = \frac{e^{\sigma\cdot(T-s)}-1}{\sigma}\,.$$

The latter expression creates problems near $\sigma = 0$ — at $\sigma = 0$, one needs to evaluate the ratio using L'Hospital's rule. To avoid such computational complexities, we evaluate the integral numerically (using trapezoidal integration with 1000 equal intervals).

Standard recent mortality tables for the U.S. imply an average male life span of 74 years and an average female life span of 80 years. We simply assume certain life spans of these amounts. There remains an issue of which life span to use in computing years left to live (i.e., $T - s$) in the numerator integral of $\kappa(s,\sigma)$. Our resolution is based on the "equivalent adults" consumption rule implicit in the Social Security System: the System's benefit formula assumes that a retired household with two adults needs 150 percent as many dollars for expenditure as a single–adult household. If the wife in our analysis is age s^f when her husband is age s, if $\phi(s,i) \equiv \min\{75 - s, 81 - s^f\}$, and if $\psi(s,i) \equiv \max\{75 - s, 81 - s^f\}$, the numerator that we ultimately employ is

$$e^{\sigma\cdot(s-50)}\cdot\Big[\frac{3}{2}\cdot\int_0^{\phi(s,i)} e^{\sigma\cdot v}\,dv + \int_{\phi(s,i)}^{\psi(s,i)} e^{\sigma\cdot v}\,dv\Big]\,. \tag{33}$$

This paper's actual regression equation is

$$\ln(NW_{is}\cdot e^{-r\cdot(s-50)}) = \beta_0 + \ln\big(Y^M_{i,50} + (1-\beta_1)\cdot Y^F_{i,50}\big) +$$
$$\beta_2\cdot(s-50) + \ln\Big(\frac{3}{2}\cdot\int_0^{\phi(s,i)} e^{\beta_2\cdot v}\,dv + \int_{\phi(s,i)}^{\psi(s,i)} e^{\beta_2\cdot v}\,dv\Big) +$$
$$\sum_{j=3}^{18}\beta_j\cdot X_{isj} + \mu_i + \epsilon_{is}\,, \tag{34}$$

where $\beta_0, ..., \beta_{18}$ are parameters to be estimated; where β_1 estimates θ and β_2 estimates σ; and where X_{isj} is a time dummy for HRS wave 1994, 1996, 1998, or 2002 (2000 being omitted since the equation has a separate constant β_0, and there are no observations from the 1992 wave in our subsample) or a time dummy for age of male retirement.[11] The

[11] In practice, our sample has no observations with male retirement age 56, and we omit one age, namely, age 62, from our list of dummy variables because our regression equation already has a constant.

median age for male retirement is about 62, and we restrict our sample to households with husbands retiring at ages 56-68. If retirement age is an exogenous variable, the latter dummies are superfluous to our estimation; conversely, they are potentially important if retirement age is a choice variable. We also restrict our sample to couples with husbands and wives no more than six years apart in age.[12]

The assumptions of our model imply that we expect an estimate, say, $\widehat{\beta_1} \in (0, 1)$. Similarly, (21) implies

$$\beta_2 = -r + \frac{r - \rho}{1 - \gamma}\,.$$

We assume $r = .0425$. The existing literature frequently sets $\rho \approx .01$ and $\gamma = 0$ to -1, and we therefore expect $\beta_2 \in [-0.01, -0.03]$.

Table 4 presents our regression results.[13] Columns 1 and 3 present nonlinear least squares regression outcomes; columns 2 and 4 present results for nonlinear feasible generalized least squares — correcting for the error correlations across observations for the same household from different survey waves. Appendix II outlines the steps for the FGLS estimation.

The outcomes should be viewed as preliminary at this stage. The estimates of theta fall in the range .26–.36. Assuming asymptotic normality, a 95 percent confidence for theta is (-0.02, 0.62) based on column 2, and it is (-0.06, 0.59) based on column 4. For sigma, a 95 percent confidence interval is (-0.04, 0.02) based on column 2, and the same based on column 4. In all cases, these overlap the set of values that we expected. The estimated intervals are, however, somewhat wide, and future work will seek to enlarge the sample (by including HRS data collected in 2004, for example) and to include additional covariates.

Discussion

Our most surprising finding is the small magnitude of our estimate of theta. Table 4 implies that as a wife enters the labor market, her household's loss of home production services is only about 30 percent as much as her market compensation; thus, the household's net gain as she finds employment outside the home is about 70 percent of what she earns at her new job.

In general, presumably changes in technology have gradually increased the marginal product of market work relative to home work. Equation (16) and our estimate $\theta = 0.30$ imply that for every one percent increase in w_t^f relative to A_t, h_t^f should increase 42.9 percent; hence, women's labor hours should have grown slowly over time. That has not been the case in practice (e.g., Manuelli *et al.* [2003]). An alternative scenario, for which some social commentators have long argued, has social and cultural norms and economic discrimination blocking women's progress in the labor market until revolutionary changes in attitudes after World War II. In that scenario, slow but persistent advances in

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[12] Recall that we model time allocation for dual adult households, where the adults have overlapping life spans.

[13] We ran OLS regressions imposing values of $\theta = 0.0, 0.1, 0.2, ..., 1.0$ and $\sigma = -0.20, -0.19, ..., 0.0, 0.01, 0.02$ and set our NLLS starting (θ, σ) to minimize the OLS sum of squared residuals.

Table 4. Nonlinear Least Squares and Nonlinear Feasible Generalized Least Squares Regression Estimates: Coefficient (Standard Deviation)

Variable	Omit Retirement Dummies		Include Retirement Dummies	
	NLLS	FGLS NLLS	NLLS	FGLS NLLS
CONSTANT	-4.9398 (0.3220)	-5.0175 (0.3832)	-4.9436 (0.3256)	-5.0761 (0.3826)
THETA	0.3546 (0.1296)	0.3023 (0.1645)	0.3277 (0.1309)	0.2658 (0.1652)
SIGMA	-0.0093 (0.0134)	-0.0065 (0.0160)	-0.0139 (0.0136)	-0.0087 (0.0161)
DUM 1994	-0.2927 (0.1697)	-0.3186 (0.1495)	-0.2774 (0.1683)	-0.3117 (0.1488)
DUM 1996	-0.1293 (0.0916)	-0.1942 (0.0779)	-0.1151 (0.0910)	-0.1884 (0.0779)
DUM 1998	-0.1178 (0.0600)	-0.1244 (0.0479)	-0.1148 (0.0595)	-0.1247 (0.0479)
DUM 2002	-0.0111 (0.0491)	-0.0134 (0.0398)	-0.0163 (0.0487)	-0.0160 (0.0398)
RET AGE 57			0.1695 (0.0839)	0.2062 (0.1043)
RET AGE 58			0.0999 (0.0945)	0.1503 (0.1161)
RET AGE 59			0.2143 (0.0866)	0.2020 (0.1018)
RET AGE 60			-0.0197 (0.0758)	-0.0048 (0.0970)
RET AGE 61			0.2170 (0.0659)	0.1952 (0.0831)
RET AGE 63			0.0678 (0.0678)	0.0490 (0.0838)
RET AGE 64			0.2768 (0.0943)	0.1562 (0.1183)
RET AGE 65			0.2724 (0.0862)	0.2658 (0.1034)
RET AGE 66			0.2487 (0.1264)	0.2457 (0.1424)
RET AGE 67			0.0385 (0.5654)	0.0348 (0.5615)
RET AGE 68			-0.0469 (0.2553)	-0.0146 (0.2921)
Addendum				
OBSERVATIONS	795	795	795	795
NO. HOUSEHOLDS	388	388	388	388
MEAN SQ ERROR	0.3243	0.3247	0.3169	0.3187
R^2	0.3274	0.3266	0.3520	0.3483

technology would lead to a build up of potential gains to female labor force participation, but the gains at first could not be realized because of social and attitudinal impediments. Once impediments dissolved, the pent up economic force was released, and women rapidly entered the labor force. Since change was in a sense overdue, net gains could be large. Our parameter estimates seem, at this point, most consistent with this type of story.

Our estimated labor supply elasticity for married women is almost exactly the same as Pencavel's [1998, table 19]. Other recent estimates tend, on the other hand, to be higher (e.g., Blundell and MaCurdy [1999]).

5. National Saving

The ratio of gross private saving in the U.S. to GDP, what we refer to as the average propensity to save, or the APS, rose from 0.17 in 1959 to 0.19 in 1979, but it fell to 0.14 in 1999. The introduction to this paper suggests that increases in the labor force participation rate for women conceivably played a role as follows: a household's saving behavior should be related to its earnings net of expenses — including expenses for replacing lost home production as women choose to work outside their home — whereas measured GDP is proportional to earnings alone; hence, the APS, which equals saving divided by GDP, may have fallen recently as growing replacement expenses caused the numerator to expand more slowly than the denominator. This section provides a quantitative assessment of this hypothesis.

With homothetic preferences, if the economy, for the sake of argument, is always in a long–run steady–state equilibrium, a household's saving at age x should be proportional to the present value of its lifetime resources.[14] This paper shows that an accurate measure of a household's lifetime resources is

$$Y^M(t) + (1-\theta) \cdot Y^F(t) \, ,$$

where $Y^M(t)$ and $Y^F(t)$ are the present value at fixed household age, say, 50, for a man and his wife if the household begins at time t. The household's saving at age x would be, say,

$$s(x) \cdot [Y^M(t) + (1-\theta) \cdot Y^F(t)] \, ,$$

with $s(x)$ depending only on age. Assuming that earnings grow with technological progress at rate g and that population grows at rate n, that $Y^M(t)$ and $Y^F(t)$ reflect average earnings within a given birth cohort, and that each household's life span is T years, we have

$$APS(t) = \frac{\int_0^T \, s(x) \cdot [Y^M(t-x) + (1-\theta) \cdot Y^F(t-x)] \cdot e^{n \cdot (t-x)} \, dx}{GDP(t)}$$

[14] Our discussion here assumes that technological progress would raise the value of time in home production, A_t in model (2), in the same proportion as it raises market wage rates over time.

$$= [Y^M(t) + (1-\theta) \cdot Y^F(t)] \cdot e^{n \cdot t} \cdot \frac{\int_0^T \; s(x) \cdot e^{-(n+g) \cdot x} \, dx}{GDP(t)} \, .$$

Thus,

$$\frac{APS(t_1)}{APS(t_0)} = \frac{\frac{[Y^M(t_1)+(1-\theta) \cdot Y^F(t_1)] \cdot e^{n \cdot t_1}}{GDP(t_1)}}{\frac{[Y^M(t_0)+(1-\theta) \cdot Y^F(t_0)] \cdot e^{n \cdot t_0}}{GDP(t_0)}} \, . \tag{35}$$

If labor earning's in total are a constant fraction of GDP over time, we have

$$\frac{GDP(t_0)}{GDP(t_1)} = \frac{[Y^M(t_0) + Y^F(t_0)] \cdot e^{n \cdot t_0}}{[Y^M(t_1) + Y^F(t_1)] \cdot e^{n \cdot t_1}} \, . \tag{36}$$

Call women's share of total earnings

$$f(t) \equiv \frac{Y^F(t)}{Y^M(t) + Y^F(t)} \, . \tag{37}$$

Together lines (35)-(37) imply

$$\frac{APS(t_1)}{APS(t_0)} = \frac{\frac{Y^M(t_1)+(1-\theta) \cdot Y^F(t_1)}{Y^M(t_1)+Y^F(t_1)}}{\frac{Y^M(t_0)+(1-\theta) \cdot Y^F(t_0)}{Y^M(t_0)+Y^F(t_0)}} = \frac{1-\theta \cdot f(t_1)}{1-\theta \cdot f(t_0)} \, . \tag{38}$$

To determine the quantitative impact of changes in $f(t)$, Table 5 compares male and female labor earnings in 1959, 1979, and 1999. We use Census micro–data from the Integrated Public Use Microdata Series (IPUMS).[15] For each year, we have a one percent sample of the U.S. population. We compute the average wage income for men and for women who are 22-62 years old — using all men and women, including those who work and those who do not (entered with zero earnings).

As one might expect, the change in female earnings relative to males is stark. In the 1960 Census, a man earned on average about four times (3.97) as much as a woman. By 2000, however, the ratio had fallen to less than two (1.80). As a share of total labor income earnings by the adult population, women account for only 20 percent in the 1960 Census, while they made up more than 35 percent in the 2000 Census.

Comparing 1959 and 1999, equation (38) then implies

$$\frac{APS(1999)}{APS(1959)} = \frac{1-\theta \cdot (.36)}{1-\theta \cdot (.20)} \, . \tag{39}$$

If $\theta = 0$, changes in female labor force participation would have had no impact upon national saving — the ratio in (39) would be 1.0. If $\theta = 1.0$ — so that each dollar of a woman's market earnings lead to a dollar less in home production — the saving rate would have fallen 19 percent because of the change in female participation. According to

[15] See Ruggles *et al.* [2004]. Note that data from the Census refers to earnings the previous year.

Table 5. IPUMS Census Male and Female Earnings Over Time: Current Dollars[a]

Census Year[b]	1960	1980	2000
Male Average Pre–Tax Wage Earnings	$4,111	$14,300	$36,253
Female Average Pre–Tax Wage Earnings	$989	$5,160	$19,554
Ratio Male to Female Earnings	3.9664	2.6307	1.7995
Ratio Male to Total Earnings	0.7986	0.7246	0.6428
Ratio Female to Total Earnings	0.2014	0.2754	0.3572
Number Males	425,022	564,381	748,358
Number Females	445,426	594,511	771,042

a. IPUMS (see text) are random samples of one percent of U.S. population. From the samples, we use all men and women ages 22 to 62, including top coded observations and zeros.

b. Data from each Census refer to earnings the year before.

Section 4, $\theta \approx 0.3$; thus, the APS should have fallen about 5 percent due to participation changes.

Using $\theta = 0.3$, increasing female labor force participation 1959-99 should, according to our illustrative calculations, have caused the APS to decline 2.5 percent; from 1979-99, the decline should have been an additional 2.5 percent. Data at the beginning of this section shows an actual 11 percent increase in the APS 1959-79 and a 30 percent decline 1979-99. Evidently the actual changes are very large relative to the model's predictions. Although the dynamics of the adjustment following changes in $f(t)$ could lead to greater short–run changes in the APS(t) than our simplified calculations imply, our low estimate of theta in Section 4 precludes a large ultimate effect.

6. Conclusion

Employing data from the Health and Retirement Study 1992-2002, as well as linked Social Security earnings records 1951-91, this paper attempts to assess the value to households of lost home production from women's entrance into the labor force. Section 4 presents our estimates. We find losses about 30 percent as large as female market earnings, implying a 70 percent net gain from women's market employment.

Our assessment of the value of lost home production is too small to explain substantial long–run declines in private saving. It has the interesting implication that recent gains in the U.S. GDP stemming from growing female labor force participation largely represent true improvements in the standard of living rather than substitution of market production for home production of the same value. Our 30 percent estimate may also shed light on the nature of the mechanisms and forces affecting the dynamic adjustment in the U.S. economy in the last 50 years toward higher labor force participation on the part of married women.

Appendix I: Lifetime Earnings

Male Earnings. As stated in the text, we assume that males work full time until they retire. We first estimate a standard earnings dynamics regression model.

For male h of age s, let y_{hs} be "real" earnings (nominal earnings deflated with the GDP consumption deflator, normalized to 1 in 1984), and let X_{hs} be a vector including a constant, a quartic polynomial in years of work experience, and time dummies. We think of the polynomial as capturing the accumulation of human capital through work experience, and we think of the time dummies as registering the impact of macroeconomic forces of technological progress. We have one additional constant, a dummy for SSA observations. To economize on parameters for the computations, we employ a linear spline for our time dummies — with constant rates of growth for 1951–60, 1961–65, 1966-70, 1971-75, etc. Our regression equation is

$$\ln(y_{hs}) = X_{hs} \cdot \beta + u_h + e_{hs} \, .$$

The regression error term has two components: an individual specific random effect u_h, and an independent, non–specific random error e_{hs}. In fact, we have one random error, e_{hs}, for observations prior to 1991 and another, $\bar{e}_{hs}$, for observations after 1991.

Let the term of the overall likelihood function for equation (1) that applies to individual h be L_h. Each individual has three types of observations: $i \in I_h$ earnings figures from the SSA history that are not right–censored; $j \in J_h$ from the SSA history that are right–censored; and $k \in K_h$ from the HRS public datasets 1992–2002 (never subject to right–censoring). Assume that u_h, e_{hs}, and $\bar{e}_{hs}$ are independent normal random variables with precisions h_u, h_e, and $h_{\bar{e}}$. Let the normal density, say, for u, be $\phi(u, h_u)$, and let the corresponding normal cumulative distribution function be $\Phi(u, h_u)$. Define

$$z_{hs} = \ln(y_{hs}) - X_{hs} \cdot \beta \, .$$

Then

$$L_h = \int_{-\infty}^{\infty} \phi(u, h_u) \cdot \prod_{i \in I_h} \phi(z_i - u, h_e) \cdot \prod_{j \in J_h} [1 - \Phi(z_j - u, h_e)] \cdot \prod_{k \in K_h} \phi(z_k - u, h_{\bar{e}}) \, du \, .$$

Our estimates of $(\beta, h_e, h_{\bar{e}}, h_u)$, which we call $(\hat{\beta}, \hat{h}_e, \hat{h}_{\bar{e}}, \hat{h}_u)$, come from

$$\min_{\beta, h_e, h_{\bar{e}}, h_u} \; - \sum_h \ln(L_h) \, .$$

For individuals with no right–censored observations, one can evaluate L_h in closed form; otherwise, we use numerical integration. We minimize the log likelihood function with

Newton's method.[1]

Earnings Data. To minimize complications from mixing full–time and part–time work, we omit observations from ages above 60 or past the man's reported retirement age. Similarly, we drop annual earnings amounts below 1500 hours $\times$ statutory minimum wage, and we drop SSA observations from years with less than four quarters of work. We assume that earnings begin at age $S = \max\{\text{years of education} + 6, 16\}$. We drop observations from ages prior to $S + 3$. Technically the HRS asks earnings and hours for the preceeding year, but we find it plausible that respondents report amounts for the current calendar year; thus, we assume survey earnings refer to 1992, 1994, etc., rather than 1991, 1993,.... As protection against coding errors, we drop earnings observations above \$1 million. As our focus is couples, we omit single men.

Male Parameter Estimates. Table A1 presents our parameter estimates. As indicated in the text, we provide separate estimates for four education groups.

Retirement. Construct the retirement age of male h, $R = R_h$, as follows: if he retires in 2002 or before and supplies the date, use it to set R; if he is retired in one wave but not retired in the preceding wave, set R equal from the date of the second wave; if he has not retired by 2002, set R from his expected retirement age; and, if dies without retiring, set R from his previously expected retirement age.[2]

Lifetime Earnings.

We predict annual earnings from our regression equation, generating $(y_S, y_{S+1}, ..., y_R)$. For ages x with SSA earnings that are not right-censored and that pass our data filters above, we substitute the actual SSA earnings figure for the imputed value of y_x. We compute annual earnings from 1992 onward by linearly interpolating and extrapolating from y_{1991} and survey data. The last step is meant to capture possible part–time work in years immediately before retirement.

We predict as follows. From our regression equation, we have

$$E[y_s \,|\, \text{data}] = e^{X_s \cdot \beta} \cdot E[e^u \,|\, \text{data}] \cdot E[e^{e_t} \,|\, \text{data}]\,.$$

Given the large sample size in our regression, we simply set $\beta = \hat{\beta}$. Since e_t is an independent random variable, assumed normally distributed, in the SSA sample we use $\sigma_e \equiv 1/\hat{h}_e$, etc., to set

$$E[e^{e_t} \,|\, \text{data}] = \begin{cases} e^{\sigma_e^2/2}, & \text{if year} \leq 1991, \\ e^{\sigma_{\tilde{e}}^2/2}, & \text{otherwise.} \end{cases}$$

[1] This paper's calculations use Compaq Visual Fortran 6.6. The minimization employs Newton's method (IMSL routine DUMIAH); we evaluate $\Phi(.)$ with IMSL function DNORDF, and we evaluate the integral for u with a 21–point Gauss–Kronrod rule (IMSL routine DQ2AGS) — truncating the bounds of integration at plus and minus six standard deviations from 0. Our version of Newton's method employs user-specified first and second derivatives. For individuals with right–censored observations, the derivatives require numerical integration.

[2] Actual analysis in Section 4 uses only surviving men who report they have retired.

The middle term in the prediction formula is the most complicated. The data provide a vector of points $\vec{z}_i$, a vector of intervals $\vec{Z}_j$ with $Z_j = [z_j, \infty)$, and a vector of points $\vec{z}_k$. Letting $f(.)$ be a density function, we have

$$f(u \mid \vec{z}_i, \vec{Z}_j, \vec{z}_k) = \frac{f(u, \vec{z}_i, \vec{Z}_j, \vec{z}_k)}{f(\vec{z}_i, \vec{Z}_j, \vec{z}_k)} \, .$$

The statistical model implies

$$f(u, \vec{z}_i, \vec{Z}_j, \vec{z}_k) = \phi(u, h_u) \cdot \prod_{i \in I} \phi(z_i - u, h_e) \cdot \prod_{j \in J} [1 - \Phi(z_j - u, h_e)] \cdot \prod_{k \in K} \phi(z_k - u, h_{\bar{e}}) \, .$$

Integrating with respect to u generates the marginal density $f(\vec{z}_i, \vec{Z}_j, \vec{z}_k)$; hence, we have

$$E[e^u \mid \text{data}] = \frac{\int_{-\infty}^{\infty} e^u \cdot \phi(u, h_u) \cdot \prod_{i \in I} \phi(z_i - u, h_e) \cdot \prod_{j \in J} [1 - \Phi(z_j - u, h_e)] \cdot \prod_{k \in K} \phi(z_k - u, h_{\bar{e}}) \, du}{\int_{-\infty}^{\infty} \phi(u, h_u) \cdot \prod_{i \in I} \phi(z_i - u, h_e) \cdot \prod_{j \in J} [1 - \Phi(z_j - u, h_e)] \cdot \prod_{k \in K} \phi(z_k - u, h_{\bar{e}}) \, du} \, .$$

Table 1 in the text summarizes the distribution of male lifetime earnings, providing present values at age 50 and in 1984 dollars.

Female Earnings. We are unwilling to assume that women necessarily work in the labor makret full time; thus, our imputation procedure is more selective than the one that we use for men.

We use the same earnings dynamics equation as for men, though for each education group we estimate the equation twice for women. First, we employ the same filters as in the case of males — calling this our "exclusive sample." For each earnings figure at the censoring limit, we use the maximum of the limit and the prediction from the first earnings dynamics equation. Table A2 presents our earnings dynamics equation coefficient estimates.[3]

Second, we enlarge our data set to what we call our "inclusive sample." This includes ages from S up to retirement, earnings below 1500 hours $\times$ statutory minimum wage, and those with less than four SSA quarters of work. In 1996 the HRS inquired about number of years of non–FICA employment and the corresponding dates. We predict non–FICA earnings for jobs prior to 1980 (after 1980, our data includes non–FICA earnings) with the second earnings dynamics equation (and the prediction steps described above). If the survey reports, for instance, two years of non–FICA work 1955-58, we impute an annual

[3] Comparing Tables A1-A2, in terms of coefficient signs and magnitudes one can see that the earnings dynamics model is much more successful for men — undoubtedly due to the heterogeneity of women's labor force hours in the birth cohort of our sample. This paper relies primarily on actual earnings from the data. Beyond that, it uses Table A1 more extensively than Tables A2-A3.

amount for each of the four years but multiply predicted values by one half. If a non–FICA prediction overlaps an observation in the data, we sum the two. Table A3 presents our second earnings dynamics equation coefficient estimates.[4]

Table 2 in the text summarizes the distribution of female lifetime earnings.

Appendix II: Feasible Generalized Least Squares Estimation[5]

At time t, let $s = s(i,t)$ be the age of household i. Define

$$e(i,t,\beta) \equiv \ln\big(NW_{i,s(i,t)} \cdot e^{-r\cdot(s(i,t)-50)}\big) - \beta_0 - \ln\big(Y^M_{i,50} + (1-\beta_1)\cdot Y^F_{i,50}\big) - \beta_2 \cdot \big(s(i,t)-50\big) - \ln\Big(\frac{3}{2}\cdot \int_0^{\phi(s(i,t),i)} e^{\beta_2\cdot v}\,dv + \int_{\phi(s(i,t),i)}^{\psi(s(i,t),i)} e^{\beta_2\cdot v}\,dv\Big) - \sum_{j=3}^{18} \beta_j \cdot X_{i,s(i,t),j}\,.$$

For each household i, we have observations for a set of times $\mathcal{T}(i)$.

Step 1. We obtain a consistent estimate of β from

$$\widehat{\beta} = \arg\max_{\beta} \sum_i \sum_{t\in\mathcal{T}(i)} e(i,t,\beta)\cdot e(i,t,\beta)\,.$$

Step 2. We generate a set of residuals $e_{it} = e(i,t,\widehat{\beta})$. Equation (34) shows the residuals represent the sum of two errors, μ_i and ϵ_{it}. We obtain a consistent estimate of the ratio of variances

$$\frac{\sigma^2_\mu}{\sigma^2_\mu + \sigma^2_\epsilon}$$

using

$$\widehat{z} \equiv \frac{\sum_i \big(\sum_{\mathcal{T}(i)} e_{it}\big)^2 - \sum_i \sum_{\mathcal{T}(i)} [e_{it}]^2}{\widehat{s^2}\cdot \sum_i \big[[t(i)]^2 - t(i)\big]}\,,$$

where

$$t(i) = \#\mathcal{T}(i)\,,$$

$$\widehat{s^2} \equiv \frac{\sum_i \sum_{\mathcal{T}(i)} [e_{it}]^2}{\sum_i t(i)}\,.$$

[4] See the preceding footnote.

[5] See, for example, Gallant [1987, ch. 2].

Step 3. For each i, let the $t(i) \times t(i)$ matrix $\Omega(i)$ have $\widehat{z}$ off of the main diagonal and 1 on the main diagonal. Let $\vec{e}(i,\beta)$ be the (column) vector $(e(i,1,\beta), ..., e(i,t(i),\beta))$. Then our FGLS estimate of β is

$$\bar{\beta} = \arg\max_{\beta} \sum_i \vec{e}(i,\beta)^T \cdot [\Omega(i)]^{-1} \cdot \vec{e}(i,\beta) \, .$$

Forming a new estimate of s^2 from the residuals based on $\bar{\beta}$, the covariance matrix for $\bar{\beta}$ is

$$\widehat{s^2} \cdot \left[\sum_i F_i^T \cdot [\Omega(i)]^{-1} \cdot F_i\right]^{-1} ,$$

where F_i is the matrix with (18–element) row $\partial e(i,t,\bar{\beta})/\partial\beta$ for each t, $t = 1, ..., t(i)$.

Table A1. Male Earnings Regression Coefficients (Standard Error), by Education; Weighted Regressions; 1984 Prices with NIPA PCE Deflator[a]

Description	Less Than High School	High School	Some College	College or more
DUM SSA	-0.0384 (0.0386)	-0.0243 (0.0250)	-0.0182 (0.0327)	0.0031 (0.0329)
CONSTANT	8.7160 (0.0578)	8.8087 (0.0418)	8.9343 (0.0642)	9.0599 (0.0835)
EXP	0.1305 (0.0101)	0.1419 (0.0073)	0.1397 (0.0102)	0.1109 (0.0135)
EXP**2/100	-0.5894 (0.0758)	-0.7606 (0.0586)	-0.7629 (0.0868)	-0.6963 (0.1188)
EXP**3/1000	0.1322 (0.0232)	0.1822 (0.0189)	0.1932 (0.0293)	0.2062 (0.0419)
EXP**4/10000	-0.0115 (0.0024)	-0.0162 (0.0021)	-0.0184 (0.0034)	-0.0233 (0.0050)
DUM 51-60	-0.0095 (0.0042)	-0.0023 (0.0032)	0.0020 (0.0055)	0.0243 (0.0080)
DUM 61-65	0.0228 (0.0046)	0.0379 (0.0034)	0.0191 (0.0053)	0.0619 (0.0068)
DUM 66-70	0.0299 (0.0048)	0.0413 (0.0035)	0.0490 (0.0054)	0.0386 (0.0069)
DUM 71-75	-0.0064 (0.0045)	-0.0037 (0.0033)	-0.0058 (0.0050)	0.0179 (0.0060)
DUM 76-80	-0.0218 (0.0042)	-0.0133 (0.0030)	-0.0245 (0.0044)	-0.0366 (0.0049)
DUM 81-85	-0.0144 (0.0041)	-0.0110 (0.0028)	-0.0029 (0.0041)	0.0169 (0.0042)
DUM 86-90	-0.0164 (0.0043)	-0.0107 (0.0029)	-0.0043 (0.0042)	0.0026 (0.0043)
DUM 91-95	-0.0037 (0.0119)	0.0103 (0.0077)	0.0130 (0.0101)	0.0343 (0.0103)
DUM 96-00	0.0372 (0.0121)	0.0324 (0.0078)	0.0338 (0.0097)	0.0249 (0.0101)
h_e	2.9986 (0.0209)	2.7966 (0.0137)	2.6460 (0.0187)	2.7476 (0.0205)
h_u	2.7660 (0.0833)	2.6664 (0.0531)	2.4399 (0.0649)	1.9540 (0.0484)
$h_{\bar{e}}$	2.4560 (0.0637)	2.3261 (0.0364)	2.3627 (0.0475)	2.1224 (0.0402)
Addendum: Summary Statistics				
Observations	15,628	35,481	17,641	18,834
Households	675	1,568	877	1,032
-ln(likelihood)	7201.3026	17616.3343	9321.1486	9106.7372

a. See text.

Table A2. Female Earnings Regression Coefficients (Standard Error), by Education; Weighted Regressions; 1984 Prices with NIPA PCE Deflator; "Exclusive Sample"[a]

Description	Less Than High School	High School	Some College	College or more
DUM SSA	-0.0741 (0.0325)	-0.0311 (0.0197)	-0.0417 (0.0278)	-0.0114 (0.0295)
CONSTANT	8.6949 (0.0547)	9.0010 (0.0361)	8.9508 (0.0706)	9.0521 (0.1564)
EXP	0.0212 (0.0112)	-0.0070 (0.0068)	0.0070 (0.0098)	-0.0762 (0.0134)
EXP**2/100	-0.1383 (0.0839)	0.0491 (0.0547)	-0.0664 (0.0817)	0.6109 (0.1176)
EXP**3/1000	0.0481 (0.0255)	0.0020 (0.0176)	0.0349 (0.0275)	-0.1713 (0.0418)
EXP**4/10000	-0.0058 (0.0027)	-0.0019 (0.0019)	-0.0056 (0.0032)	0.0156 (0.0051)
DUM 51-60	0.0149 (0.0048)	0.0058 (0.0034)	0.0134 (0.0074)	0.0426 (0.0162)
DUM 61-65	0.0136 (0.0054)	0.0134 (0.0036)	0.0292 (0.0059)	0.0389 (0.0097)
DUM 66-70	0.0266 (0.0048)	0.0243 (0.0033)	0.0129 (0.0049)	0.0392 (0.0073)
DUM 71-75	0.0017 (0.0042)	-0.0036 (0.0029)	0.0047 (0.0042)	-0.0035 (0.0058)
DUM 76-80	0.0024 (0.0038)	0.0075 (0.0026)	0.0192 (0.0037)	-0.0001 (0.0047)
DUM 81-85	0.0027 (0.0035)	0.0078 (0.0024)	0.0188 (0.0033)	0.0298 (0.0041)
DUM 86-90	0.0004 (0.0036)	-0.0016 (0.0024)	0.0157 (0.0033)	0.0229 (0.0040)
DUM 91-95	-0.0004 (0.0100)	0.0099 (0.0060)	0.0224 (0.0084)	0.0377 (0.0090)
DUM 96-00	0.0320 (0.0095)	0.0229 (0.0053)	0.0253 (0.0071)	0.0396 (0.0077)
h_e	3.9150 (0.0318)	3.4870 (0.0172)	3.3102 (0.0228)	3.1271 (0.0263)
h_u	3.1901 (0.0934)	2.8766 (0.0526)	2.7064 (0.0697)	2.1436 (0.0599)
$h_{\bar{e}}$	2.6811 (0.0677)	2.7892 (0.0406)	2.5524 (0.0469)	2.6801 (0.0547)
Addendum: Summary Statistics				
Observations	9,223	26,015	13,899	10,609
Households	798	1,946	999	782
-ln(likelihood)	1927.8852	7996.1526	5141.4456	4603.7158

a. See text.

Table A3. Female Earnings Regression Coefficients (Standard Error), by Education; Weighted Regressions; 1984 Prices with NIPA PCE Deflator; "Inclusive Sample"[a]

Description	Less Than High School	High School	Some College	College or more
DUM SSA	-0.0814 (0.0591)	-0.0236 (0.0339)	-0.0653 (0.0472)	-0.0415 (0.0503)
CONSTANT	7.5107 (0.0832)	8.2312 (0.0504)	8.2383 (0.0862)	7.8950 (0.1583)
EXP	0.0923 (0.0129)	0.0251 (0.0071)	-0.0350 (0.0098)	-0.0512 (0.0135)
EXP**2/100	-0.3493 (0.0978)	-0.1496 (0.0574)	0.1984 (0.0804)	0.4207 (0.1197)
EXP**3/1000	0.0799 (0.0300)	0.0697 (0.0185)	-0.0389 (0.0266)	-0.0830 (0.0423)
EXP**4/10000	-0.0076 (0.0031)	-0.0097 (0.0020)	0.0018 (0.0029)	0.0024 (0.0050)
DUM 51-60	-0.0317 (0.0083)	-0.0176 (0.0051)	0.0225 (0.0088)	0.0788 (0.0173)
DUM 61-65	0.0232 (0.0104)	0.0082 (0.0065)	0.0218 (0.0097)	0.0121 (0.0145)
DUM 66-70	0.0343 (0.0102)	0.0222 (0.0067)	0.0387 (0.0093)	0.0319 (0.0125)
DUM 71-75	-0.0056 (0.0096)	0.0138 (0.0062)	0.0325 (0.0087)	0.0196 (0.0113)
DUM 76-80	0.0095 (0.0091)	0.0258 (0.0058)	0.0517 (0.0081)	0.0438 (0.0101)
DUM 81-85	-0.0215 (0.0089)	-0.0052 (0.0056)	0.0358 (0.0076)	0.0262 (0.0092)
DUM 86-90	0.0184 (0.0090)	0.0100 (0.0056)	0.0269 (0.0076)	0.0392 (0.0090)
DUM 91-95	0.0145 (0.0175)	0.0213 (0.0098)	0.0373 (0.0135)	0.0308 (0.0144)
DUM 96-00	0.0350 (0.0139)	0.0298 (0.0078)	0.0163 (0.0103)	0.0431 (0.0111)
h_e	1.1161 (0.0065)	1.1375 (0.0043)	1.1160 (0.0061)	1.0546 (0.0072)
h_u	1.3572 (0.0364)	1.4392 (0.0252)	1.4521 (0.0365)	1.2222 (0.0333)
$h_{\bar{e}}$	1.3197 (0.0239)	1.4644 (0.0164)	1.4047 (0.0208)	1.5002 (0.0257)
Addendum: Summary Statistics				
Observations	17,375	43,648	21,685	15,615
Households	981	2,252	1,120	862
-ln(likelihood)	23,826.32	57,277.90	28,878.51	20,661.63

a. See text.

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