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Does Stockholding Provide Perfect Risk Sharing?

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Abstract

This paper analyzes the extent of risk-sharing among stockholders. Wealthy households play a crucial role in many economic problems due to the substantial concentration of wealth and asset holdings in the U.S. data. Hence, to evaluate the empirical importance of market incompleteness, it is essential to determine if idiosyncratic shocks are important for the wealthy who also have better insurance opportunities compared to the average household. We study a dynamic structural model where each period households compare the benefits of stockholding with a per-period trading cost and decide whether to participate in the stock market. Due to the endogenous entry decision, the testable implications of perfect risk-sharing take the form of a sample selection model. To eliminate the selection bias, we implement a semiparametric estimation method recently proposed by Kyriazidou (2001). Using data from PSID we strongly reject perfect risk-sharing for stockholders, but perhaps surprisingly, find no evidence against it among non-stockholders. The results are robust to a number of changes in the test method, such as including future wages into the instrument set, and testing from long time differences. We offer some explanations based on private information problems and the resulting idiosyncratic production risk borne by wealthy households. Finally, we strongly reject risk-sharing for the whole population consistent with existing literature. These findings indicate that, if anything, market incompleteness may be more important for the wealthy, and suggest further focus on risk factors that primarily affect this group, such as business risks.

Keywords: Perfect risk-sharing, incomplete markets, semiparametric estimation, Generalized Method of Moments, limited stock market participation.

JEL Classification Codes: C33, G11, D52.

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1 Introduction

In the last few years models with incomplete markets and uninsurable idiosyncratic shocks have achieved a central place in many fields of economics. These models are used to study a wide range of economic problems, such as wealth inequality, asset prices, business cycle dynamics, fiscal policy and so on. For example, in Aiyagari (1994)-style models wealth inequality results from the desire of households to accumulate a buffer stock wealth to self-insure against adverse shocks in the absence of complete markets (Krusell and Smith (1998); and Castaneda, et. al. (2002)). Similarly, there is a growing sense that uninsurable income shocks play an important role in determining asset prices by making households more reluctant to take financial risk, causing them to demand higher returns (among others, Constantinides and Duffie (1996); Heaton and Lucas (1996); Storesletten, et. al (2001), Brav, et. al. (2002)).

A major motivation for this literature has been the decisive empirical rejection of perfect risk-sharing—the hypothesis that individuals can insure against all idiosyncratic shocks and are thus able to equate the growth rate of their marginal utilities to one another. A number of empirical studies have found households' consumption growth (or more precisely their marginal utility growth) to be correlated with certain idiosyncratic shocks—and income shocks in particular—violating the premise of perfect insurance (Cochrane (1991); Nelson (1994); Townsend (1994); Attanasio and Davis (1996); and Hayashi, Altonji and Kotlikoff (1996)). If, on the other hand, perfect risk-sharing (PRS) could be attained, idiosyncratic risk would be vanquished and play no role.

An important point to note is that these studies test if PRS holds *for the whole population.* However, given that asset holdings and wealth are extremely concentrated—basically 90 percent of non-housing wealth and 95 percent of equity is held by the top 20 percent of the U.S. population—wealthy households play a crucial role in many economic interactions. Thus, for a satisfactory analysis of many of the issues mentioned above, it is essential to determine the extent of risk-sharing *among wealthy households.* For example, the main source of idiosyncratic uncertainty in most heterogenous-agent models is labor income risk which is estimated to be substantial by empirical studies on labor earnings.[1] Nevertheless, if the wealthy are able to diversify this risk effectively, modeling these shocks as "uninsurable" may overstate the amount of risk in the economy.

On the one hand, there are good reasons to suspect that wealthy households may stand a better chance of achieving perfect risk-sharing compared to the rest of the population. After all, the top 20 percent of the population almost exclusively trade in stock markets, which is arguably the most sophisticated market-based risk sharing mechanism. Moreover, with ample assets that can be used as collateral, borrowing constraints are less likely to be an obstacle to optimal portfolio formation. Finally, empirical evidence indicates that high-skilled individuals face lower unemployment risk as

[1] See, among others, Moffit and Gottschalk (1995), Meghir and Pistaferri (2001), Gourinchas and Parker (2002), and Storesletten, et. al (2002).

well as smaller labor income shocks conditional on being employed (c.f. Kydland (1984); Gourinchas and Parker (2002) and the references therein). Putting these pieces together, it seems possible that the empirical rejection of PRS for the whole population may be driven by lack of insurance among the poor and may not provide a justification that idiosyncratic shocks are important for the small minority of wealthy.

On the other hand, wealthy households are exposed to production risk to a much larger degree compared to the rest of the population. For example, private capital—which is roughly as large as the capital in publicly traded companies, and is potentially difficult to insure due to private information problems—is concentrated among the wealthy, exposing them to entrepreneurial income risk not faced by other households.[2] These differences in the kinds of risks faced and in the access to financial markets suggest that the severity of market incompleteness may be quite different for each group. Thus, in this paper we formally investigate the extent of risk-sharing among the wealthy (stockholders) and among the rest of the population (non-stockholders).

A second motivation for studying the risk insurance role of financial markets is normative. In all modern economies enormous amounts of public funds are dedicated to providing social insurance in the form of unemployment insurance, welfare programs, and so on. However, the canonical model of portfolio decision—going all the way back to Arrow (1964)—attributes a central role to financial assets in the optimal sharing of risk: if all households have access to a complete set of securities the resulting allocations will be Pareto optimal leaving no role for government intervention. Therefore it is compelling to ask if, empirically, financial markets are able to provide risk insurance effectively, in which case government policy could better serve by encouraging participation in these markets. This can be accomplished, for example, by favorable tax treatment of investment income (eliminating dividend taxation?), or by raising public awareness as was done by the British government in the 1980s, which resulted in stock market participation rates almost tripling in just a few years.

The interpretation of the risk-sharing tests faces a difficulty common to all classical hypothesis testing: with a sufficiently large sample size all empirical hypothesis will be rejected. At the same time, lack of rejection with a fixed sample size or a given instrument set may indicate the low power of the test. Thus, a potentially more useful way to approach this question is to view perfect insurance as an ideal benchmark and to assess the extent of risk-sharing among stockholders *compared to* the rest of the population by their distance from this benchmark. That is, with a given sample size and instrument set, if one rejects risk-sharing for one group but not for the other, this difference can be interpreted as differences in the extent of insurance among each group. Hence, we ask if the wealthy (stockholders) are able to share risk more effectively than the rest (non-stockholders), where the latter serves as a control group.

[2] Gentry and Hubbard (1998), Heaton and Lucas (2000), and Moskowitz and Vissing-Jorgensen (2001) document the extreme concentration of private capital and the business risks faced by entrepreneurs.

We employ a flexible preference specification which allows for non-separabilities between consumption and the leisure times of head and spouse, and incorporates household-specific preference shifters. Moreover, ignoring preference heterogeneity can lead to false rejections of the PRS hypothesis (Ogaki and Zhang (2001)), so we allow stockholders and non-stockholders to have different risk aversion and leisure elasticity parameters.

We consider an economy with a full set of financial securities which are traded in the stock market. Thus all shocks are *potentially* insurable. Each period households decide whether to participate in the stock market by paying a one-time entry cost and a per-period participation cost, or to stay out and invest in a single risk-free asset. These fixed costs are intended to capture the disutility associated with learning how to invest, and once in the market, the time spent monitoring one's portfolio. Households also make optimal portfolio and labor supply decisions.

Due to the endogenous stock market entry decision, the testable implications of the risk-sharing hypothesis for stockholders take the form of a dynamic sample selection model (Tobit Type II, in the terminology of Amemiya (1985)) where the participation decision rule serves as the selection equation. To eliminate the selection bias, we implement a semiparametric GMM estimator recently proposed by Kyriazidou (2001) for panel data models, which does not require strong distributional assumptions about the error terms. To our knowledge, this is the first implementation of this estimator.

Our findings can be summarized as follows. Using data from the Panel Study of Income Dynamics (PSID) on U.S. households we strongly reject perfect risk-sharing among stockholders, but perhaps surprisingly, we find *no* evidence against it among non-stockholders with consistently high p-values across experiments. This result is robust to a number of changes, such as including future values of wages into the instrument set (as advocated by Hayashi et. al (1996)), using long time differences of the moment conditions (as suggested by Attanasio and Davis (1996)), using the implications of PRS for the marginal utility of leisure, and so on. Finally, we strongly reject risk sharing for the whole population consistent with existing literature, suggesting that the rejections reported in earlier studies are likely to be due to the failure of insurance not among the poor, but instead, among the wealthy. Therefore, if anything, incomplete markets and idiosyncratic shocks are more important for wealthy households. This finding in turn underscores the importance of focusing on risks primarily faced by the wealthy such as entrepreneurial income risk, which are recently being incorporated into incomplete markets models (c.f., Cagetti and Denardi (2001); Angeletos and Calvet (2002); Chari et. al. (2002); and Smith and Wang (2002)).

These results are consistent with a number of recent findings in the literature. For example, Brav, Constantinides and Geczy (2002) show that even after accounting for limited stock market participation, market incompleteness among stockholders play a critical role in explaining the equity premium puzzle. Second, Attanasio and Davis (1996)—while rejecting perfect insurance for the entire population—also find that it fails most significantly for the highest educated individuals whereas there is much weaker evidence against it among lower educated groups (p. 1247, fig 3). Since education and wealth are positively correlated (Table 1) their finding suggests the same pattern of risk sharing in the population as the one uncovered here (more on this in Section 7.1).

In fact, imperfect risk-sharing among the wealthy can be viewed as the (constrained) efficient outcome in an environment with private information and long-term contracts. Private information naturally arises in entrepreneurial activities (largely undertaken by the wealthy), or may result from the interaction between managers and firm owners. As is well-known, with private information the social planner will implement *in*complete risk-sharing to induce proper incentives (Rogerson (1985), Ligon (1998), among others). We discuss this point further in Section 8.

Finally, by allowing for heterogeneity in preferences, we can address another interesting question. Starting with Mankiw and Zeldes (1991) researchers have documented various differences between the choices of stockholders and non-stockholders, such as their consumption processes, wealth levels, etc. A natural question is whether these differences mainly reflect different investment opportunity sets faced by these groups, or whether they also reveal more fundamental heterogeneity in preferences. In our framework, we incorporate the differences in budget sets, and still find statistically significant heterogeneity in the curvature of consumption and the labor supply elasticities of the two groups.

In terms of method and approach this paper is most closely related to a number of studies which test for perfect insurance among smaller groups in the population. The discouraging rejection of PRS in the whole population (Cochrane (1991); Nelson (1994); Attanasio and Davis (1996); Hayashi et. al (1996)), led researchers to focus on smaller units who have strong ties with the hope of uncovering full insurance within these groups. Examples include households living in the same geographical regions (Hess and Shin, 2000), inhabitants of small villages in various underdeveloped countries (Townsend (1994); Udry (1994); Ligon (1998); and Ogaki and Zhang (2001)), and finally, family members (Hayashi, Altonji and Kotlikoff (1996)).

The paper is organized as follows. In the next section we set up the model. We specify the parameterization for the empirical work in Section 3. In Section 4 we develop the econometric techniques to analyze this problem in the presence of selection bias. Then in Section 5 we describe the data, and in Section 6 we discuss the estimation of the selection equation and the construction of kernel weights. Section 7 explains how we carry out the estimation and presents the results for the tests of risk-sharing with various instrument sets. We then discuss some explanations for the findings, and conclude in Section 8.

2 The Model

We consider an economy which runs for $T < \infty$ periods. There are a finite number of households, each with a life span of $\tau < T$ periods, and each is composed of at least a head (henceforth called husband) and a spouse. The uncertainty structure in this economy is treated as a probability space $(\Omega, \mathcal{F}, P)$ where each element in Ω denotes a particular realization of all random variables in the economy for all dates. The information available to households can be represented as a sequence of increasing $\sigma-$ algebras (information filtration) $(\mathcal{F}_1 \subseteq \mathcal{F}_2 \subseteq ... \subseteq \mathcal{F}_T)$ and E_t denotes the mathematical expectation conditional on $\mathcal{F}_t$.

Households derive utility from consumption as well as from husband's and spouse's leisure times. To capture heterogeneity in the population, we assume that each household's utility is also influenced by household-specific preference shifters, summarized in the vector $\mathbf{Z}_{nt}$. Specifically, the intertemporal preferences of household n is given by

$$E_0 \left[\sum_{t=0}^{\tau} \beta^t u \left(C_{nt}, L_{1nt}, L_{2nt}, \mathbf{Z}_{nt} \right) \right], \tag{1}$$

where C_{nt} denotes consumption in period t; L_{1nt} and L_{2nt} are the leisure times of head and spouse respectively. The period utility function, u, is continuously differentiable and concave in the choice variables for each value of $\mathbf{Z}_{nt}$.

Financial Markets and Participation

Let s_t denote a date-event pair (state), which constitutes a complete description of uncertainty *for all the economy* that is realized in t, and let $S^t = (s_1, s_2, ..., s_t)$ be the history of all states realized up through period t. For example, s_t will contain the realization of wages of all households, the return on all assets in the economy, etc. Further, let s_{nt} denote the component of state relevant for the optimization problem of household n, such as the history of its own wages, $\{W_{1n\kappa}, W_{2n\kappa}\}_{\kappa=1}^{t}$, the return on all assets, and so on, and S_n^t contain its relevant history.

Each node s_t branches out into S possible states (successor nodes) in the next period. There is a complete set (S) of Arrow securities (one-period contingent claims) at every state, each paying one unit of consumption good in exactly one state of the world tomorrow. From this description, it is clear that all shocks are potentially insurable. Moreover, in each period a risk-free bond is also available for investment in this economy.

The risk-free bond is traded among all households freely (that is, without incurring any fixed or proportional transaction costs) but the same does not hold for other assets. It is plausible to think that it takes time and effort to initially learn how financial markets work (a one-time cost), and once in the market, to monitor the performance of one's portfolio (a fixed cost per-period).[3]

[3] In the empirical part, we look at the U.S. data from 1984 to 1992, when these informational costs were presumably

Empirical evidence suggests that both of these costs are likely to be significant in practice.[4] To capture this idea, we assume that households have to pay a one-time cost of Ψ^0 to enter the stock market, and a fixed cost of Ψ_n^P in every period they participate thereafter.[5] In principal, we can also allow both these costs to change through time at the aggregate level. As will become clear in the empirical part where we entertain this possibility, this modification does not affect the main results, so we assume them to be fixed through time to save on notation. Note that we have not introduced *proportional* transactions costs, because this would preclude perfect risk sharing even among stock market participants.

As for the income of households, both the husband and the spouse are endowed with one unit of time in each period which they allocate between working in the market and leisure time spent at home. Provided that the labor market is competitive, the wage rate of each individual in the economy can be written as follows:

$$W_{int} = \delta_w \left(\widehat{\mathbf{Z}}_{nt} \right) \overline{W_t}, \tag{2}$$

where $\delta_w(\widehat{\mathbf{Z}}_{nt})$ is an efficiency index function; $\widehat{\mathbf{Z}}_{nt}$ is a vector of household characteristics possibly containing some elements not included in $\mathbf{Z}_{nt}$, and $\overline{W_t}$ is the market wage rate. Although we do not need to make any assumptions about the labor market for tests of risk sharing, this wage equation will be useful in some specifications as an additional moment condition to increase asymptotic efficiency of the estimator.

To better understand the choices facing a typical household, it is useful to express the decision problem recursively. Let d_n^0 be an indicator function which takes on the value 1 if the household has ever participated in the stock market before and is zero otherwise. Each period households decide whether to participate in the stock market in the current period by paying the fixed cost, $\Psi_n^P + \left(1 - d_n^0\right) \Psi^0$, or to stay outside and trade in the risk-free bond only. Define $\mathbf{q}(s) \equiv (q_1, q_2, ..., q_S)$ to be the price vector of the Arrow securities when the state is s, and q_0 to be the bond price. Similarly, let $\mathbf{k}_n \equiv (k_{1n}, k_{2n}, ..., k_{Sn})$ denote a current stockholder's portfolio choice vector of Arrow securities, and k_{0n} be the bond holdings of a non-stockholder. We drop the time subscript, and

higher than what they are now. Of course, these costs fell substantially in the past few years which also boosted participation in financial markets. This relationship is also implied by the model described here.

[4] Vissing-Jorgensen (2000) estimates these costs from micro data and conclude that even modest fixed costs are sufficient to keep a large fraction of households out of the stock market. Luttmer (1999) and Paiella (2000) also estimate participation costs, although they do not distinguish between the different types.

[5] The one time entry cost, Ψ^0, is not essential for any of the results in the paper; it is introduced to show the generality of the test results even under state dependence. However, allowing Ψ^0 to be individual-specific would substantially complicate the estimation of the participation decision, so we avoid it.

denote next period's variables by primes. Then a household's problem is:

$$\upsilon\left(\omega_n, d_n^0; S_n, \mathbf{Z}_n\right) = \max_{d_n \in \{st, no\}} \left[\upsilon^{st}\left(\omega_n, d_n^0; S_n, \mathbf{Z}_n\right), \upsilon^{no}\left(\omega_n, d_n^0; S_n, \mathbf{Z}_n\right)\right]$$

where

$$\begin{aligned}
\upsilon^{st}\left(\omega_n, d_n^0; S_n, \mathbf{Z}_n\right) &= \max_{C_n, L_1, L_2, \mathbf{k}_n} \left\{u\left(C_n, L_{1n}, L_{2n}, \mathbf{Z}_n\right) + \beta E_t\left(\upsilon\left(\omega_n', d_n^{0\prime}; S_n', \mathbf{Z}_n'\right)\right)\right\} \\
& \quad s.t \\
C_n + \mathbf{q}^T(s)\,\mathbf{k}_n &\leq \omega_n + \sum_{i=1}^{2}(1 - L_{in})W_{in}(s) - \Psi_n^P - \left(1 - d_n^0\right)\Psi^0 \\
\omega_n'\left(s'\right) &= k_{s'n}.1,
\end{aligned}$$

and,

$$\begin{aligned}
\upsilon^{no}\left(\omega_n, d_n^0; S_n, \mathbf{Z}_n\right) &= \max_{C_n, L_1, L_2, k_{0n}} \left\{u\left(C_n, L_{1n}, L_{2n}, \mathbf{Z}_n\right) + \beta E_t\left(\upsilon\left(\omega_n', d_n^{0\prime}; S_n', \mathbf{Z}_n'\right)\right)\right\} \\
& \quad s.t \\
C_n + q_0(s)\,k_{0n} &\leq \omega_n + \sum_{i=1}^{2}(1 - L_i)W_{in}(s) \\
\omega_n' &= k_{0n},
\end{aligned}$$

where υ^{st} and υ^{no} are the value functions of current stockholders and non-stockholders respectively, and ω_n denotes financial wealth.

Note the difference between the budget sets of current stockholders and non-stockholders. In particular, the former group chooses an unrestricted $S \times 1$ portfolio vector implying that they can transfer any (budget-feasible) amount of wealth to a particular state in the next period. Thus markets are dynamically complete within the stock market community. In contrast, non-stockholders are restricted to choosing a state-independent (constant) wealth level, k_0, for the next period.

2.1 Perfect risk-sharing

The model that we laid out in the previous section features a choice in every period between a complete markets world and an incomplete markets world and thus has a more complicated structure than the models underlying previous tests of risk-sharing. To fix ideas, let us first look at the canonical complete markets model, where all households can trade in a complete set of securities in all periods. In this case, each household maximizes

$$E_0\left[\sum_{t=0}^{\tau} \beta^t u\left(C_{nt}, L_{1nt}, L_{2nt}, \mathbf{Z}_{nt}\right)\right]$$

subject to a single lifetime budget constraint:

$$E_0\left[\sum_{t=0}^{\tau}\beta^t\lambda(s_t)(C_{nt}-\sum_{i=1}^{2}(1-L_{int})W_{int})\right]\leq 0, \tag{3}$$

where $\lambda(s_t)$ is the time-zero price of one unit of consumption in state s_t (i.e., the state-price density). If, on the other hand, households have only access to a single risk-free bond—markets are incomplete—then they face a sequential budget constraint:

$$C_{nt}+q_t k_{0n,t+1}\leq k_{0,t}+\sum_{i=1}^{2}(1-L_{int})W_{int}, \qquad 0\leq t\leq\tau.$$

For the incomplete markets case, the first order conditions are:

$$\begin{aligned} u_1\left(C_{nt},L_{1nt},L_{2nt},\mathbf{Z}_{nt}\right) &= \mu_{nt}, \\ u_i(C_{nt},L_{1nt},L_{2nt},\mathbf{Z}_{nt}) &\geq \mu_{nt}W_{int}, \qquad i=2,3, \end{aligned} \tag{4}$$

where μ_{nt} is the Lagrange multiplier on the time t budget constraint, and u_i is the partial derivative of u with respect to its ith argument.

On the other hand, in the complete markets case, if we let ϕ_n be the multiplier associated with the lifetime budget constraint (3), it can be easily shown that (4) holds true but with some extra structure imposed:

$$\frac{1}{\phi_n}u_1\left(C_{nt},L_{1nt},L_{2nt},\mathbf{Z}_{nt}\right) = \lambda_t, \tag{5}$$

$$\frac{1}{\phi_n}u_i(C_{nt},L_{1nt},L_{2nt},\mathbf{Z}_{nt}) \geq \lambda_t W_{int}, \qquad i=2,3. \tag{6}$$

Taking the ratio of (5) for periods t and $t+1$ to eliminate the unobservable component, ϕ_n, we get

$$\frac{u_1\left(C_{nt+1},L_{1nt+1},L_{2nt+1},\mathbf{Z}_{nt+1}\right)}{u_1\left(C_{nt},L_{1nt},L_{2nt},\mathbf{Z}_{nt}\right)}=\frac{\lambda_{t+1}}{\lambda_t}. \tag{7}$$

This last equation provides a clear illustration of perfect risk sharing: the marginal utility growth of all households (on the left hand side) should only be a function of aggregate variables (the right hand side). An alternative way to phrase this statement is that after accounting for preference shifters and the leisure times of head and spouse, the cross-sectional distribution of consumption growth should not be correlated with the cross-section of any household-level variable. Notice how radical this prediction is: it requires (not ex-ante, but) ex-post marginal utility growth to be equated across households. Any deviation from this equality that may be observed in data can only be attributed to measurement error or unobserved heterogeneity (through Z_{nt}). On the other hand, with incomplete markets, shocks to marginal utilities, μ_{nt}, are individual specific and are potentially different for every agent. The main test of PRS is to estimate the relationship in (7) and then see if any idiosyncratic variable is correlated with the resulting error term.

Turning to the model with endogenous participation presented in the previous section, there is the additional complication that households optimally enter and exit the stock market in different states and thus face complete markets and incomplete markets at different points in time. Fortunately, in any given period, the same risk-sharing condition still holds for households who are in the stock market in that period. We derive this result in Appendix A. Thus we can use (7) *together* with the participation decision (selection equation) to test for risk-sharing among stockholders.

2.2 The participation decision

In general a closed form solution to the participation choice is not available, although it is easy to see that a numerical solution can be obtained by solving backwards starting from the last period of a household's life (for example, as in Keane and Wolpin (1997)). The focus of this paper is on risk-sharing, and our interest in the participation decision is mainly for having a good specification of households who self-select into the stock market. Thus, rather than explicitly solving for d_{nt}, we will seek variables that determine the participation decision, which we can obtain from the optimization problem above.

A household enters the market if $\upsilon^{st}\left(\omega_n, d_n^0; S_n, \mathbf{Z}_n\right) > \upsilon^{no}\left(\omega_n, d_n^0; S_n, \mathbf{Z}_n\right)$ and stays outside if the reverse holds which means the decision rule will be a function of the vector of state variables: $\left(\omega_{nt}, d_{nt}^0, S_n^t, \mathbf{Z}_{nt}\right)$. In the context of this particular portfolio choice problem, S_{nt} contains the household's wage history, $\mathbf{W}_n^t = \{W_{1n\kappa}, W_{2n\kappa}\}_{\kappa=1}^t$, and asset prices. Furthermore, we assume that asset returns are serially independent, so only the vector $\mathbf{W}_n^t$ enters the decision rule.[6]

The cost parameter Ψ_n^P will affect participation through the budget constraint of υ^{st}. Similarly, if $d_{nt}^0 = 0$, then Ψ^0 enters the decision rule. Other parameters of the model will also play a role in this decision, but they are all assumed identical across the population except those already summarized in $\mathbf{Z}_{nt}$, so they can be soaked up into the functional form. To sum up, we can write the binary decision rule for a typical agent as

$$d_{nt} = 1\left\{\pi\left(\omega_{nt}, d_{nt}^0, \mathbf{W}_n^t, \mathbf{Z}_{nt}, \Psi_n^P, \Psi^0\right) \geq 0\right\},$$

[6] The independence assumption on returns, aside from being quite common in the literature, is made for the following reason: As we discuss in the next section, to deal with selection bias, we will assume that the error terms in the selection equation are *i.i.d.* This latter is a common assumption in the dynamic discrete choice literature (c.f., Keane and Wolpin (1997); Honoré and Kyriazidou (1999); Chintagunta, et. al. (2001)) and is a necessary assumption for the particular estimation strategy that we will use. See Section 4. Independence of returns is a clean way to satisfy this assumption without necessitating a measurement of returns faced by each agent.

where $1\{\cdot\}$ is an indicator function, and $\pi(\cdot)$ is determined by the solution to the problem.

Notice that the set of variables included in $\pi(\cdot)$ represent all *potential* determinants of stockholding; it is likely that empirically only a subset of them are significant factors in participation choice. For example, variables which affect υ^{st} and υ^{no} symmetrically will leave $\upsilon^{st} - \upsilon^{no}$ unchanged and will have no impact on participation. Thus, identifying the significant determinants of stockholding is ultimately an empirical question, which we address in Section 6.1

3 Empirical Investigation

Since perfect risk-sharing imposes restrictions on marginal utilities, the specification of preferences is especially important for the purpose of this paper. A false rejection of PRS may very well result from ignoring heterogeneity, non-separabilities, and so on. For example, Ogaki and Zhang (2001) strongly reject PRS when they restrict the risk aversion parameter across the population, but find no evidence against it once they allow for preference heterogeneity.

We assume the following period utility function for each group

$$U^i(C, L_1, L_2, \mathbf{Z}) \equiv \delta_0(\mathbf{Z}) C^{\rho_0^i} L_2^{\rho_2} + \delta_1(\mathbf{Z}) L_1^{\rho_1^i} L_2^{\rho_3},$$

for $i = st, no$, and where $\rho_0^{st} = \rho_0^{no} + a_0$, and $\rho_1^{st} = \rho_1^{no} + a_1$.[7]

This specification for preferences is quite flexible, and is in fact, more general than most of those considered in the previous literature. The first sub-utility can be interpreted as a Cobb-Douglas home-production function where (food) consumption serves as capital and female leisure hours as labor input. One can view this specification as the first two sub-utilities of a more general function in which non-food consumption enters in a separable manner. For further discussion and justification of this separability (which has also been the maintained assumption in all previous papers using PSID) see Section 5.

The second sub-utility captures the possible non-separability between the leisure time by head and spouse. Non-separable specifications in both sub-utilities have empirical support (Browning and Meghir (1991), Altug and Miller (1990)). Another possibility is to have male leisure also enter the first sub-utility. Hayashi, et. al. (1996) test for this possibility and do not find any support for it. Moreover, if in fact male leisure and consumption are non-separable, tests based on equation (7) are invalid due to observational equivalence.[8]

Some components of the vector $\mathbf{Z}_{nt}$ may not be observable to the econometrician. Hence, it is convenient to write $\mathbf{Z}_{nt} = (\mathbf{x}_{nt}, \varepsilon_{1n}, \varepsilon_{2n}, \varepsilon_{1nt}, \varepsilon_{2nt})$, where $\mathbf{x}_{nt}$ is a vector which represents the

[7] This specification restricts the female leisure parameters ρ_2 and ρ_3 to be the same across two groups. This is not dictated by economic theory but is rather an identifying assumption. In Section 7.2, we investigate the robustness of our results to relaxing this restriction.

[8] See Attanasio and Davis (1996, page 1235) for a discussion of this point.

observable component, and the εs denote the unobservables. Each sub-utility is weighted by indices which are log-linear functions of $\mathbf{Z}_{nt}$:

$$\delta_m^i(\mathbf{Z}_{nt}) = \frac{1}{\rho_m^i} \exp\left(\mathbf{B}_m \mathbf{x}_{nt} + \varepsilon_{mn} + \varepsilon_{mnt}\right), \quad m = 0, 1.$$

Here $\mathbf{B}_m$ is a fixed vector of coefficients; ε_{mn} represents the fixed household-effect, and ε_{mnt} is a zero-mean disturbance term which varies both over time and across households. Further assumptions on the error terms will be stated in the next section. Note that each subutility is scaled by ρ_0 and ρ_1.

Finally we want to parameterize the participation equation. One empirical issue that arises is that d_{nt}^0 may not be observable given the short time dimension of available panel data sets. For example, in PSID if a household did not own stocks in 1984 when stockholding information was collected for the first time, then there is no way to determine the value of d_{nt}^0. A reasonable way to get around this problem is to assume that the knowledge acquired by paying this fixed cost depreciates quickly (in one period) if a stockholder leaves the stock market. In this case what matters is whether a household was in the stock market in the previous period or not, and $d_{n,t-1}$ replaces d_{nt}^0 in the selection equation.

For tractability, we specialize the selection function $\pi(\cdot)$ to a linear form, which allows us to write the decision rule as a standard binary choice equation. Substituting the observable and unobservable parts of $\mathbf{Z}_{nt}$,we obtain

$$d_{nt} = 1\{\phi d_{n,t-1} + \boldsymbol{\theta}\boldsymbol{\varphi}_{nt} + \eta_n - \eta_{nt} \geq 0\}, \tag{8}$$

where ϕ is a scalar related to Ψ^0; $\boldsymbol{\theta}$ is a fixed vector of coefficients; $\boldsymbol{\varphi}_{nt} \equiv \left(\omega_{nt}, \mathbf{W}_n^t, \mathbf{x}_{nt}\right)$; $\eta_n \equiv \Psi_n^P + \varepsilon_{0n} + \varepsilon_{1n}$, and $\eta_{nt} \equiv -\left(\varepsilon_{0nt} + \varepsilon_{1nt}\right)$.

4 Methodology

Under the null hypothesis that the model described in Section 2 is correct, the equilibrium allocation of household n is a single realization of the random vector $\{C_{nt}, L_{1nt}, L_{2nt}\}$ which satisfies the first order conditions with probability one.

Using the parameterization for preferences, the risk-sharing condition for stockholders (7) yields our first moment condition. Define $\Delta\nu_{1nt} \equiv -\Delta\varepsilon_{0nt}$, where Δ_t denotes the difference operator between t and $t-1$. After taking logarithms, first differencing, and rearranging, we obtain

$$-\rho_2 \Delta l_{2nt} = -\Delta \ln(\lambda_t) + \mathbf{B}_0 \Delta \mathbf{x}_{nt} + (\rho_0 + a_0 - 1)\Delta c_{1nt} + \Delta\nu_{1nt}, \tag{9}$$

where lower-case letters denote the natural logarithms of their upper-case counterparts (except for

$\mathbf{x}_{nt}$).

Although this equation by itself is sufficient to test for risk-sharing, it cannot identify all the parameters of interest. For that purpose, we add another moment condition which is valid for all households (so that both groups' preferences are identified). Dividing the first equation by the second in (4) and (5), we get the same equation for both groups:

$$\frac{U_2\left(C_{nt}, L_{1nt}, L_{2nt}, \mathbf{Z}_{nt}\right)}{U_1\left(C_{nt}, L_{1nt}, L_{2nt}, \mathbf{Z}_{nt}\right)} = W_{1nt}, \tag{10}$$

which is the familiar intra-temporal efficiency condition equating the marginal rate of substitution between consumption and leisure to the wage rate.[9] This equation can be used as a basis for estimation. Given the parametrization for preferences, again it is convenient to take logarithms, then difference (10), and define $\Delta\nu_{2nt} \equiv \Delta\varepsilon_{1nt} - \Delta\varepsilon_{0nt}$. (The reason for time-differencing this equation will become clear shortly.) Then we can write

$$\begin{aligned}(\rho_2 - \rho_3)\Delta l_{2nt} &= -\Delta w_{1nt} + (\mathbf{B}_1 - \mathbf{B}_0)\Delta \mathbf{x}_{nt} - (\rho_0 - 1)\Delta c_{1nt} \\ &\quad +(\rho_1 - 1)\Delta l_{1nt} - a_0(d_{nt}\Delta c_{1nt}) + a_1(d_{nt}\Delta l_{1nt}) + \Delta\nu_{2nt}.\end{aligned} \tag{11}$$

4.1 Sample selection bias

Since $\Delta\nu_{1nt}$ and $\Delta\nu_{2nt}$ have zero mean by construction, it might seem reasonable to look at (9) and (11) as defining orthogonality conditions $E\left[\Delta\nu_{int}\right] = 0$, $i = 1, 2$, which could then be estimated using GMM. However, this strategy is not directly applicable in this framework due to sample selection.

To clearly see this point, let us first consider the PRS condition (9) for stockholders. Also, for the sake of discussion set $\phi \equiv 0$, so that there is no state dependence. Since under the null hypothesis only stockholders are able to share risk perfectly, the appropriate moment condition is:

$$E\left[\Delta\nu_{1nt} \mid d_{nt}d_{n,t-1} = 1\right] = 0. \tag{12}$$

Even if we assume that the vector $\{\nu_{1nt}, \eta_{nt}\}$ is serially independent conditional on the regressors

[9] Notice that the MRS equation holds only when the head is working in a given period, which can potentially cause another selection problem. But, first, there is evidence in the literature that male unemployment is largely involuntary (see Ham (1986), and the discussion on page 55 of Pencavel (1986), for example). Second, and may be more convincingly, Altug and Miller (1990) estimate a Tobit specification for selection into the labor market and find that the error term in the selection equation has a small and insignificant correlation with the error in the MRS equation. Moreover, we eliminate far fewer households compared to Altug and Miller since we only require the head to work for two consecutive years to be included in the estimation (whereas they require this for all fourteen years), so this problem should be even less serious in this case.

in the selection equation, for (12) to hold we still need

$$
\begin{aligned}
E\left[\nu_{1nt} \mid d_{nt}=1\right] &= E\left[\nu_{1nt-1} \mid d_{n,t-1}=1\right] \Rightarrow \\
E\left[\nu_{1nt} \mid \eta_{nt} \leq \boldsymbol{\theta}\boldsymbol{\varphi}_{nt}+\eta_n\right] &= E\left[\nu_{1nt-1} \mid \eta_{nt-1} \leq \boldsymbol{\theta}\boldsymbol{\varphi}_{nt-1}+\eta_n\right].
\end{aligned}
$$

First, in general both sides of this equation will be non-zero, because η_{nt} is correlated with ν_{1nt}: the unobservable preference shock, ε_{0nt}, is included both in η_{nt} and in ν_{1nt}, naturally creating a correlation. In other words, unobservable preference shocks will affect the risk-sharing condition as well as the participation decision.

Of course, it is still possible that these conditional expectations equal each other and the *difference* is zero. But this is not likely to be the case either since these expectations are functions of $\boldsymbol{\theta}\boldsymbol{\varphi}_{nt}$, and will vary over time as this selection index changes. Thus, time differencing will not eliminate the selection bias term which is likely to be time-varying.

A similar selection problem manifests itself in the estimation of the MRS condition (11) because our sample selection procedure described in next section eliminates households who change their stockholding status during the sample period: $d_{nt} \neq d_{n,t-1}$.[10] Even though this moment condition holds for the whole population unlike the PRS condition, the error term $\Delta\nu_{2nt}$ has zero mean over the entire population, whereas we need this expectation to be zero over the sample *that we observe*: $E\left[\Delta\nu_{2nt} \mid d_{nt}=d_{n,t-1}\right]=0$. The rest of this section will detail the method that we will use to consistently estimate this model also allowing for state dependence ($\phi \neq 0$).

We employ the semiparametric fixed-effects estimator recently proposed by Kyriazidou (2001) adapted to deal with endogenous regressors. She considers the case where all regressors in the main equation are either strictly exogenous or lagged endogenous variables. On the other hand, in our case consumption and leisure are likely to be correlated with the contemporaneous preference shocks captured in ν_{1nt}. Second, measurement error can also induce correlation between regressors and error terms. As long as these measurement errors are multiplicative, taking the logarithm of the first order conditions will push them into the error term ν_{1nt}, and we can still consistently estimate the model by instrumenting for these regressors.

Let y_{nt} be a vector of instruments. We make the following two assumptions.

Assumption A1. $\{(\nu_{nt}, \eta_{nt})\}_{t=1}^{\tau}$ is *i.i.d* over time for all n, conditional on $\zeta_n \equiv \{\boldsymbol{\varphi}_n, \eta_n, y_{n0}, d_{n0}\}$, where $\boldsymbol{\varphi}_n \equiv (\boldsymbol{\varphi}_{n1}, ..., \boldsymbol{\varphi}_{n\tau})$.

Assumption A2. (ν_{nt}, η_{nt}) is independent of y_{ns} for all $s < t$, and for all n conditional on ζ_n.

The first assumption is the same as condition $(A1')$ in Kyriazidou (2001). The second one is a slight weakening of her assumption $(A2')$ that allows us to have endogenous variables in the main

[10] Of course identification in the selection equation comes from observations on switchers, otherwise ϕ would not be identified. So, we start with a sample including switchers, after estimating the selection equation switchers are eliminated for the main equation to be consistently estimated.

equation and to instrument for them using lagged dependent variables.

The idea behind this estimator is in the spirit of Powell (1987) who proposed pairwise comparisons in the *cross-section* to eliminate the sample selection bias. In our context, it can be explained as follows. For the sake of discussion, let us continue to assume $\phi \equiv 0$. From the discussion above, it is clear that the term $E\left[\nu_{1nt} \mid \eta_{nt} \leq \boldsymbol{\theta}\boldsymbol{\varphi}_{nt} + \eta_n, \zeta_n\right]$ will remain fixed if and only if the selection index, $\boldsymbol{\theta}\boldsymbol{\varphi}_{nt}$, is constant in two consecutive periods. In this case, time differencing eliminates not only the fixed-effect for that household but also the selection bias term. Thus, we replace (12) with

$$E\left[\Delta\nu_{1nt} \mid \Delta\boldsymbol{\theta}\boldsymbol{\varphi}_{nt} = 0, d_{nt}d_{n,t-1} = 1, \zeta_n\right] = 0. \tag{13}$$

However, one immediately observes that if $\boldsymbol{\varphi}_{nt}$ contains any continuous variables, the set of households $\{\boldsymbol{\theta}\boldsymbol{\varphi}_{nt} = \boldsymbol{\theta}\boldsymbol{\varphi}_{nt-1}\}$ may be very small, or even empty. One strategy is then to assign a weight to each observation which is inversely proportional to the change in the index for that household, $\Delta\boldsymbol{\theta}\boldsymbol{\varphi}_{nt}$, such that asymptotically only observations with constant indices are included in the estimation.

By following a similar argument one can show that in the presence of state dependence $(\phi \neq 0)$,

$$E\left[\Delta\nu_{1nt} \mid \Delta\boldsymbol{\theta}\boldsymbol{\varphi}_{nt} + \phi\left(1 - d_{n,t-2}\right) = 0,\ d_{nt}d_{n,t-1} = 1, \zeta_n\right] = 0. \tag{14}$$

Note that when $d_{n,t-2} = 1$ this conditioning set reduces to the same one above. Thus, if a household is observed for three consecutive periods as a stockholder, the same kernel-weighted estimator will deliver consistent estimates in the presence of state dependence. Similarly, for the MRS equation:

$$E\left[\Delta\nu_{2nt} \mid \Delta\boldsymbol{\theta}\boldsymbol{\varphi}_{nt} = 0, d_{nt} = d_{n,t-1} = d_{n,t-2}, \zeta_n\right] = 0. \tag{15}$$

Observe that this moment condition holds for households in both groups as long as they do not change their stockholding statuses for three consecutive periods.

Implementation

In the first step, the dynamic discrete choice model is consistently estimated to obtain an estimate of $\boldsymbol{\theta}$ (called $\widehat{\boldsymbol{\theta}}$). Then, using this estimate of the selection index, we construct weights which we take to be "kernel density" functions of the following form:

$$\widehat{\psi}_{nt}^{N} = \frac{1}{h_N} K\left(\frac{\Delta\widehat{\boldsymbol{\theta}}\boldsymbol{\varphi}_{nt}}{h_N}\right),$$

where $K(\cdot)$ is a scalar function which satisfies certain smoothness and regularity conditions,[11] and

[11] Basically, we assume that $\int |K(x)|\, dx < \infty$, $\int K(x)\, dx = 1$, and consider symmetric kernels: $\int xK(x)\, dx = 0$.

h_N is a sequence of "bandwidths" which tends to zero as the sample size $N \to \infty$. For a fixed magnitude of difference, $\Delta\widehat{\boldsymbol{\theta}}\boldsymbol{\varphi}_{nt}$, the weight $\widehat{\psi}_{nt}^{N}$ shrinks as N increases, while for fixed N, a larger deviation in the index corresponds to a smaller weight.

Let $\mathbf{f}(\boldsymbol{\alpha})$ denote a column vector of orthogonality conditions that is satisfied in the population, $\mathbf{f}(\boldsymbol{\alpha}, n)$ be its sample counterpart for the nth observation, and $\boldsymbol{\alpha}$ be the vector of identifiable parameters in that system. Further, let Φ_N be a stochastic matrix that converges in probability to a finite non-stochastic limit Φ_0. The estimator of $\boldsymbol{\alpha}$ is

$$\widehat{\alpha}_N = \arg\min \left\{ \mathbf{G}_N(\boldsymbol{\alpha})' \, \Phi_N' \Phi_N \mathbf{G}_N(\boldsymbol{\alpha}) \right\},$$

where $\mathbf{G}_N(\boldsymbol{\alpha}) \equiv \frac{1}{N}\sum_{n=1}^{N} \widehat{\psi}_{nt}^{N} \mathbf{f}(\boldsymbol{\alpha},n)$ and a prime denotes the transpose of a matrix. Essentially what we do is to stack the sample counterparts of each moment condition for observation n into the vector $\mathbf{f}(\boldsymbol{\alpha},n)$ and weight each element with the kernel weights and sum it across households to obtain $\mathbf{G}_N(\alpha)$. This estimator is consistent and asymptotically normal with $\sqrt{Nh_N}$ convergence rate (Kyriazidou, 2001).[12]

In order to construct $\mathbf{f}(\boldsymbol{\alpha},n)$, for each year t we pick $(r_j \times 1)$ dimensional vectors of instruments y_{jnt} satisfying

$$E\left[y_{jnt}\Delta\nu_{jnt} | \Delta\boldsymbol{\theta}\boldsymbol{\varphi}_{nt} = 0, d_{nt} = d_{n,t-1} = d_{n,t-2}\right] = 0, \tag{16}$$

for each $j = 1, ..., J$, where each subscript j labels a different disturbance term. For each year then we have $\left(\sum_{j=1}^{J} r_j\right)$ moment conditions for estimation denoted by $\mathbf{f}_t(\boldsymbol{\alpha}, n)$. The indexing of r_j makes it clear that a different set of instruments may be interacted with each error term, but the set of instruments are fixed across time. We reduce the panel data estimation into cross-section by forming the following $T^*\left(\sum_{j=1}^{J} r_j\right)$ dimensional vector where T^* is our panel length

$$\mathbf{f}(\boldsymbol{\alpha}, n) = \left(\mathbf{f}_1(\boldsymbol{\alpha}, n), ..., \mathbf{f}_{T^*}(\boldsymbol{\alpha}, n)\right)'. \tag{17}$$

5 The Data

This section briefly discusses the data for our empirical work.[13] Appendix B explains our sample selection criteria as well as variable definitions and construction in more detail.

Moreover the smoothness of the kernel affects the asymptotic convergence rate which imposes restrictions on the empirical choice of the function $K(\cdot)$. We will work with a Gaussian kernel which satisfies these conditions.

[12] In the absence of a formula for the optimal weighting matrix for this kernel-weighted GMM estimator, we will choose Φ_0^* such that: $\Phi_0^{*T}\Phi_0^* = E\left((\widehat{\psi}_{nt}^{N})^2 \mathbf{f}(\boldsymbol{\alpha},n)^T \mathbf{f}(\boldsymbol{\alpha},n)\right)^{-1}$ which is optimal in the standard GMM case.

[13] The sample used for estimation in this paper as well as the codes necessary to replicate all the reported results are available at www.econ.rochester.edu/guvenen/RSH2002.htm.

Table 1: A List of Key Variables and Their Simple Statistics

	Stockholders	*Non-stockholders*	*All*
Hours and Earnings			
(i)Average annual	2213	2177	2189
hours of husband	(646.1)	(686.5)	(672.1)
(ii)Average annual	1451	1501	1483
hours of spouse	(741.5)	(706.8)	(718.2)
(iii)Average hourly	$17.83	$10.41	$12.99
earnings of husband	(13.79)	(7.45)	(9.76)
(iv)Average hourly	$10.10	$6.82	$7.96
earnings of spouse	(9.08)	(5.71)	(6.88)
(v)Average annual	$5249	$4419	$4708
Food consumption	(2806)	(2253)	(2445)
Demographic Variables			
(i)Average age of	43.8	39.9	41.2
husband	(11.3)	(11.4)	(11.3)
(ii)Average Education	6.07	4.9	5.31
of Head	(1.54)	(1.62)	(1.59)
(iii)Average household	3.3	3.6	3.5
size	(1.13)	(1.21)	(1.18)
(i) Number of Observations	3178	5763	8941
(ii)Percentage of estimation sample	34.8%	65.2%	100%

Note: Standard Deviations are in paranthesis. All the statistics reported are for the final estimation sample as described in the Data Appendix.

We use data from the Panel Study of Income Dynamics (PSID), which has been extensively used in the literature to study risk-sharing. We choose married or permanently cohabiting couples as the basic unit of our economy. Although currently PSID final release data is available from 1968 to 1993, we start our sample from 1984, which is the first time data on stockholding was collected.

Let t stand for year $(t + 1980)$. For each household $n \in \{1, ..., N\}$ we have data on (a) annual leisure hours for head and spouse, denoted L_{1nt} and L_{2nt} for $t \in \{1, ...12\}$, respectively; (b) real average hourly earnings of head and spouse, denoted W_{1nt} and W_{2nt} for $t \in \{1, ...12\}$, respectively; (c) age of head and spouse, denoted A_{1nt} and A_{2nt} for $t \in \{1, ...12\}$, respectively; (d) real household food consumption expenditures (which is the sum of "food at home," "food away from home," and "the cash value of food stamps"), C_{nt}, for $t \in \{1, .., 6, 9, ...12\}$; (e) number of household members, a_{nt}, for $t \in \{1, ...12\}$; (f) completed education of head, E_{nt}, for $t \in \{1, ...12\}$; (g) a dummy indicating whether the household is a stockholder, d_{nt}, for $t \in \{4, 9\}$. Table 1 provides the summary statistics of the data for both stockholders and non-stockholders.

Briefly, we select a family into the sample in year t if the head: (i) was in the study for four consecutive years including 1984 or 1989; (ii) was married to the same spouse in t and $t-1$; and (iii) had positive labor hours in t and $t-1$. For each four-consecutive-year period, we use the first two years to construct instruments and the last two years for estimation. Furthermore, we follow the standard procedure of eliminating households who had missing data on some key variables listed in Appendix B. Filtering out these observations leaves a total of 8941 household-years (observations) that can be used in estimation.

The definition of stockholding adopted in this paper includes ownership of shares of stock in publicly held corporations, mutual funds, investment trusts, including stocks in IRA's. All households who indicate they do not own any of these assets are considered non-stockholders for that year.[14]

PSID collects stock ownership data every five years (from 1984 on) whereas for our empirical work we need this information for every year. Thus, we identify a household as a stockholder (alternatively, non-stockholder) in every year between 1984 to 1989, if he is present in the sample in both years as a stockholder (non-stockholder). Second, if a household switches between these two groups from 1984 to 1989, we eliminate those observations from the sample between these two dates since we are not able to determine when the switch exactly happens. Clearly, this step creates another selection bias, which the econometric method is able to handle as explained in the previous section (see equation 15). Finally, for years after 1989 we take the status of a household as it is given in 1989. This identification scheme is not a perfect one, but notice that the estimation method asymptotically assigns zero weight to an observation if the probability of being a stockholder changes ($\Delta\boldsymbol{\theta}\boldsymbol{\varphi}_{nt} \neq 0$). Thus, a household who moves in, or who moves out of the stock market between 1984 and 1989 will receive a small (and asymptotically zero) weight in estimation since this move is likely to be accompanied by a change in the selection index (i.e., the probability of participation in the stock market). Given that in the data stockholding is a rather persistent phenomenon, this identification should provide a reasonable approximation.

Before closing this section, there are a few points concerning the use of food data for consumption that should be addressed. First, separability between food and non-food has been the maintained assumption in all studies on risk-sharing using PSID data mentioned above, which makes our results comparable (Altug and Miller (1990); Cochrane (1991); Hayashi, et. al. (1996); Hess and Shin (2000), etc.). Second, Atkeson and Ogaki (1996) provide evidence that food and non-food consumption are separable.[15] Moreover, Ogaki and Zhang (2001, 2002) find virtually the same

[14] This definition does not include indirect ownership of stocks, for example, through pension funds. First, in 1980s indirect holding was much more modest; defined contribution plans became much more popular in the last decade. Second, we may not want to include these indirect stock owners in our sample anyway, because we want portfolio holdings to represent optimal choices on the part of households. Pension plans are not likely to satisfy this condition and may not be suitable for sharing risk unless one is willing to pay steep withdrawal penalties.

[15] See also, however, Attanasio and Browing (1995) who argue against separability between food and non-food consumption.

results regarding PRS when they replicate their tests using non-durable consumption insead of food expenditures, which is reassuring. Third, a possible concern could be that food consumption may not be sufficiently variable causing risk-sharing tests to have low power. But, if anything, the volatility of food consumption (from PSID) is higher than the volatility of non-durables consumption calculated from the Consumption Expenditure Survey. This is true for both stockholders and non-stockholders. Finally, many of the papers mentioned above are able to reject PRS for the different groups that they analyze which suggests that the power may not be low after all.

6 Estimation

6.1 First step: The selection equation

The first step in the procedure is to consistently estimate the parameter $\boldsymbol{\theta}$ from the participation equation in order to determine which of the potential factors are empirically significant for the stockholding decision. This question has received a lot of attention recently, especially with the boom in stock market participation during the 1990s. Researchers have estimated static and dynamic discrete choice models of participation in financial markets. Using different data sets and increasingly more general approaches, each subsequent paper has, by and large, confirmed the findings of earlier ones.

Among these, Haliassos and Bertaut (1995) estimate a static logit model for stockholding using individual-level data from the Survey of Consumer Finances with eighteen explanatory variables which can be grouped as follows:

1. *Demographic variables*: age, age square, sex.
2. *Socio-Economic variables*: marital status, race, education level (less than high school, high-school, some college).
3. *Preferences*: attitude toward risk (low, intermediate), willingness to give up liquidity.
4. *Income and wealth*: labor income (square root), financial net worth (cubic root), whether majority of wealth is inherited.
5. *Occupation* (whether managerial occupation).

Essentially, this list contains all the variables included in our specification of the selection equation (8) except for lagged participation (and labor income is used instead of wages). They find that the following variables are significant at 10% level: (1) race, (2) less than high school education, (3) high school education, (4) risk aversion measures, (5) managerial occupation, (6) labor income, and (7) financial net worth.

Extending this work, Hurst, Luoh, and Stafford (1998) and Vissing-Jorgensen (2000) analyze the *dynamics* of stockholding decision in PSID data using a probit estimation. In consensus, they find the same variables above to be significant determinants (at 1% or lower) of the decision to become a stockholder. The latter paper investigates true state dependence using 1984 stockholding status as d_n^0 and finds strong evidence supporting it (ϕ is significant at 0.1% significance level). Moreover, unlike Haliassos and Bertaut (1995), these studies use the time average of recent labor income (typically past three to five years) rather than current income for $\mathbf{W}_n^t$, which is closer to our specification.

An important point to observe about these findings is that except for labor income and financial wealth (6 and 7 above), these explanatory variables represent individual (or household) characteristics that do not change over time. Since kernel weights are constructed based on the *time change* in the selection index, all but two of these regressors become redundant.[16] Consequently, the coefficient estimates on these fixed characteristics do not affect the second step estimation. The only information we need to know is which of the regressors are significant, and those above include all variables that are significant at 10%, to be conservative. The conclusion we draw from this analysis is that, in order to correct for the selectivity bias, we need to mainly consider movements in labor income and financial wealth through time.

The fact that these studies use labor income instead of wages is not critical. First, in our sample total family labor wages—which we define as $\widehat{w}_{nt} \equiv \sum_{\kappa=t-2}^{t} \log(W_{1n\kappa} + W_{2n\kappa})$ analogous to Hurst et. al (1998) and Vissing-Jorgensen (2000)—has a correlation of 0.86 with the square root of family labor income, suggesting that it can be a good proxy for that variable. Second, average labor hours are very similar across stockholders and non-stockholders, so income variable is on average a scaled version of wages. Nevertheless, labor income is an endogenous variable (due to labor hours) and is thus likely to be correlated with the preference shifters included in $\Delta\nu_{1nt}$, which is not allowed by Assumption A1 above. In our work then, we substitute $\widehat{w}_{nt}$, for labor income.[17]

A similar argument applies to financial wealth which may potentially be correlated with all past values of ν_{1nt}. To eliminate this endogeneity, we first observe that the average wealth level of a household over the relatively short sample period can be captured in the fixed effect η_n, which is allowed to have arbitrary correlation with ν_{1nt} for all t. Persistent differences in wealth levels *across* households are probably more significant in determining participation than smaller year-to-year variations over time *for a given* household. Second, year-to-year changes in household's financial

[16] Note that even though we have not explicitly stated in the selection equation, $\pi(\cdot)$, in Section 2.2, heterogeneity in preference parameters is also likely to affect the participation decision. The papers mentioned here all include risk aversion measures, so they control for differences in ρ_0. But again, given that preferences are fixed over time for a given household, they will not affect the kernel weights.

[17] In fact, using alternative proxies for labor income, for example, by assigning fixed weights corresponding to average male and female labor hours of each household, or even using the labor income itself (ignoring the correlation) yields qualitatively the same results as those reported here.

wealth is likely to be correlated with changes in labor income. This relationship is clear when there is a shock to labor income and agents respond by saving or dissaving. But another possibility is that households may experience shocks directly to wealth, such as a bad entrepreneurial investment. The wage variable in PSID also includes the labor portions of business and farm income as well as of trade, gardening, and so on (Appendix B gives the exact definition) suggesting that some part of these shocks will also be captured in labor income. Thus, the proxy used above for labor income will also capture some of the short-term variation in financial wealth during the sample period. To sum up, the two variables, $\widehat{w}_{nt}$ and η_n, are likely to account for a sizeable part of the variation in the selection index that is due to variations in income and wealth.[18]

Note that none of the studies mentioned above find age to be a significant determinant of stock market participation. Although participation does increase with age—from around 20 percent at age 25 up to 35 percent at age 50 in our sample period—that increase is fully explained by the life cycle evolution of labor income and wealth. Guiso, Haliassos and Japelli (2002) confirm this finding when they estimate the participation decision in 6 European countries and use age dummies to allow for possible nonlinearities. If there was age dependence that part of the selection index would not difference out completely and would have to be taken into account in kernel construction. While this is feasible, the non-dependence makes implementation easier.

Finally, if there was not extensive empirical evidence on the determinants of participation decision, or if the coefficient estimates were crucial for the estimation of the main equation, then it would be desirable to re-estimate the selection equation using less restrictive (non-parametric) methods. One could then implement the semiparametric estimator recently proposed by Honoré and Lewbel (2002), as long as the selection equation contains one regressor which is independent of both the fixed effect and the disturbance term.

6.2 Second step: The main equation

In this section we discuss the estimation and testing of the main equation. The main equation that we are interested in testing is the PRS condition for stockholders (9) reproduced below for

[18] Note that we are using the time average of wages, $\widehat{w}_{nt}$, not the contamporaneous value, to construct kernel weights so that households who receive wage shocks in the current period do not necessarily receive a smaller weight. As a specific example, an individual who works at time $t-3$, then loses his job and remains unemployed for two years, and finds a new job at t for a comparable wage would receive a weight close to 1 even though he experienced a large shock at time t. So, even asymptotically, there will be a positive measure of households who experience income shocks in the current period which is important for testing for PRS. Moreover, the cross-sectional standard deviation of $\Delta\widehat{w}_{nt}$ for stockholders and non-stockholders are very close to each other (0.36 versus 0.38) implying that on average both groups receives similar kernel weights. Second, households face many other shocks that are not captured by income movements such as health shocks, strikes, and involuntary moves (Cochrane 1991; Hayashi et. al 1996). We are also testing for insurance against these other sources of shocks which are captured by including instruments for leisure hours of head and spouse, past consumption levels, and so on.

convenience:

$$-\rho_2 \Delta l_{2nt} = -\Delta \ln(\lambda_t) + \mathbf{B}_0 \Delta \mathbf{x}_{nt} + (\rho_0 + a_0 - 1)\Delta c_{1nt} + \Delta \nu_{1nt}. \tag{9'}$$

As noted earlier, the MRS equation (11) will be used to identify the remaining parameters of the model:

$$\begin{aligned}(\rho_2 - \rho_3)\Delta l_{2nt} \quad = \quad & -\Delta w_{1nt} + (\mathbf{B}_1 - \mathbf{B}_0)\Delta \mathbf{x}_{nt} - (\rho_0 - 1)\Delta c_{1nt} + (\rho_1 - 1)\Delta l_{1nt} \\ & -a_0(d_{nt}\Delta c_{1nt}) + a_1(d_{nt}\Delta l_{1nt}) + \Delta \nu_{2nt}.\end{aligned} \tag{11'}$$

Another benefit of this MRS equation is that it can yield some insight into the validity of the instruments that we use for testing full insurance. Since all the variables in the PRS equation (except λ_t) also appear in the MRS condition, any measurement error in the former equation will also show up in the latter. The same is true for the unobserved part of family characteristics. Thus, if an instrument is correlated with $\Delta \nu_{1nt}$, it will also be correlated with $\Delta \nu_{2nt}$. So, when we estimate the MRS equation alone, it will get rejected indicating a potential problem with that instrument. We will then avoid those instruments when testing PRS.

Finally, the following "seemingly unrelated" wage equation (2) can further increase the asymptotic efficiency of the estimator through the correlation of error terms:

$$\Delta w_{1nt} = \Delta \ln(\overline{W_t}) + \mathbf{B}_2 \Delta \widehat{\mathbf{x}}_{nt} + \Delta \nu_{3nt}. \tag{18}$$

As for the empirical specification of $\mathbf{x}_{nt}$, we choose a square and a cubic polynomial of husband's age, A_{1nt}^2, A_{1nt}^3, and the household size, H_{nt}.[19] These variables are intended to capture the changing household structure and needs through the life-cycle (such as consumption requirements, spouse's time to care for children, etc.).[20]

There are a number of different ways the risk-sharing hypothesis can be tested. The first and most obvious one is to estimate (9') alone for stockholders and use Hansen's $J - test$ as a model specification test. If stockholders are not sharing risk perfectly, then marginal utility growth cannot be completely explained by aggregate shocks and the resulting error term will be correlated with idiosyncratic variables. By including household-level variables in the instrument set this correlation will be caught by the $J - test$ as a model specification error. This idea also forms the basis of the previous tests implemented in the literature.

[19] Similarly, $\widehat{\mathbf{x}}_{nt}$ includes a constructed experience variable: $\left(A_{1nt} + E_{nt}^2\right)$ where E_{nt} is the categorical education variable.

[20] We have also experimented with adding a linear term in husband's age, spouse's age, husband's education level, and the number of children in the household, but these additions left the hypothesis testing results essentially unchanged.

A second method, whose advantage will become clear in a moment, is the following: First estimate the MRS condition (18) which holds for the entire population. Then append (9') multiplied by d_{nt} as an additional moment condition and estimate the two jointly, and test for PRS as an overidentifying restriction of the model. Specifically, if the additional orthogonality condition imposes p_1 extra restrictions and identifies p_2 additional parameters (and $p_1 > p_2$) then Nh_N (effective sample size) times the increment in the GMM criterion function is distributed χ^2 with $(p_1 - p_2)$ degrees of freedom.[21] This test has the flavor of a Chow test, in the sense that we test a restriction for a subsample of the population.

The second approach has the advantage of exploiting more information thereby increasing the efficiency of the estimator and the power of our hypothesis test. This latter method is our preferred test, but for completeness we will also report test results with the former one.

6.2.1 How to choose the kernel bandwidth, h_N?

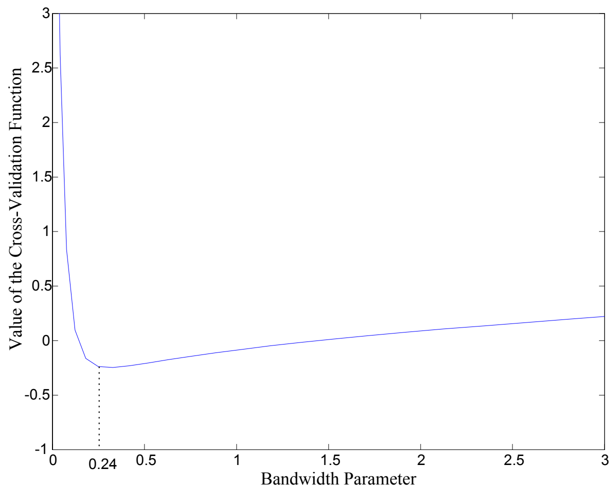
A standard Gaussian density is used for the kernel function $K(\cdot)$ which is asymptotically nearly optimal (Epanechnikov, 1969). Because asymptotically optimal kernel functions perform only slightly better even in the limit, normal density is a reasonable choice in practice.

The first step is to choose h_N. As is usually the case with semiparametric methods, asymptotically optimal methods for selecting the bandwidth provide little guidance for practical implementation with a fixed sample size. However, observing that the estimated weighting function, $\widehat{\psi}_{nt}^{N}$, has a structure similar to a kernel density estimator, a sensible approach is to select h_N as the cross-validated value for the estimation of the density of the selection index, $\boldsymbol{\theta\varphi}_{nt}$ (see, Chen and Khan 2002 for a similar application of this idea). Hence, the bandwidth is chosen by minimizing the mean integrated squared error of the kernel density estimator as described in Fan and Gijbels (1996, page 150). This procedure yields $h_N^* = 0.24$.

Figure 1 plots the criterion function which is quite flat between 0.2 to 0.5, although it increases steeply outside this region. Due to the exponential nature of weights, small differences in the value of h_N in this range results in large variations in kernel weights. For example, a household whose selection index changes by 50% between two periods is weighted by 0.61, 0.13 and 0.004 for $h_N = 0.5$, 0.24 and 0.15. To make sure that our conclusion is robust to values of h_N in this range, we will also report the results throughout for $h_N = 0.5$ as well.

[21] See Eichenbaum, Hansen and Singleton (1988) for derivation in the standard GMM case; it is straightforward to derive it for the estimator here using the results in Kyriazidou (1997, 2001). However, in the latter case, because the estimator is asymptotically biased, the distribution of the test statistic is non-central chi-squared, with the non-centrality parameter (NCP) equal to the squared mean of $\left(1/\left(\sqrt{Nh_n}\right)\right)\sum_{n=1}^{N}\widehat{\psi}_N f_j(\alpha_j, n)$. Even though this quantity can be estimated in principle, this is very difficult in practice (see Bierens 1987 for a detailed discussion). The Monte Carlo experiments in Kyriazidou (1997, 2001) suggest that this bias is very small in general, which implies that the NCP is also small. We use the central chi-squared distribution to perform the hypothesis tests. In the worst case, this will bias the results towards rejection if the NCP is large (see Davidson and MacKinnon 1993, pp. 412-414).

Figure 1: Cross-Validation Objective as a Function of h_N



7 Results

In this section we report our empirical findings. First, we investigate the effect of the bias correction method introduced in Section 4.1. Second, we discuss the results obtained from tests of risk-sharing for stockholders, non-stockholders and the whole population. Then we check the robustness of the results: we repeat the tests of PRS using long time differences of the moment conditions (as advocated by Attanasio and Davis (1996)), using future values of wages as instruments (as argued by Hayashi et. al (1996)), by reducing the number of instruments (to check sensitivity to small sample performance) and by using the moment conditions implied by the cross-sectional distribution of leisure. We conclude by discussing parameter estimates.

In PSID consumption (food) data is not available in 1987 and 1988, which leaves us with six time differences that can be used in estimation: 83-84, 84-85, 85-86, 89-90, 90-91, 91-92.

Instruments. Our main set includes the following nine variables for the estimation using the time difference between $t-1$ and t: a constant; age of head at time t, A_{1nt}; age of spouse at time t, A_{2nt}; household size at time t and $t-1$, H_{nt} and H_{nt-1}; change in log consumption from year $t-3$ to $t-2$, $\Delta c_{n,t-2}$; change in log wage of husband from year $t-3$ to $t-2$, $\Delta w_{1n,t-2}$; change in log spouse's wage from year $t-3$ to $t-2$, $\Delta w_{2n,t-2}$; a dummy indicating whether household is a stockholder or not, d_{nt}.

This instrument set is used for the MRS condition. Notice that we have not included the first lags of variables which are susceptible to measurement error, such as consumption and wages, because the resulting correlation with variables in the MRS equation would make them invalid. For the PRS condition, we add the contemporaneous wage growth of the head, $\Delta w_{1n,t}$, as well as its first lag, $\Delta w_{1n,t-1}$, to increase the power of the risk sharing test. Since this equation does not contain head's wage, possible measurement error in wages is not likely to cause a problem. Finally, in order to keep the total number of instruments small, we exclude the stockholding dummy, d_{nt}, and use change in household size instead of levels (ΔH_{nt-1}) . Finally, for the wage equation, (18), that will be added later we exclude female wage change, $\Delta w_{2n,t-2}$, and consumption growth, $\Delta c_{n,t-2}$, but add the education of head, E_{1nt} to the instrument set above.

7.1 Tests of risk-sharing

As explained earlier, we test for risk-sharing among stockholders, and to provide a benchmark, also among non-stockholders. In order to investigate the effect of bias correction, we first estimate the model by a naive GMM estimator which ignores the sample selection problem. First, column 1A of Table 2, displays the results for the estimation of the MRS equation only. The χ^2 statistic for model specification test has a value of 46.5 with 43 degrees of freedom, so the MRS condition cannot be rejected at conventional significance levels ($p - value = 32.9\%$). Next, we append the PRS condition for stockholders (1B). The incremental value in the χ^2 from this extra moment condition is 81.1 (127.6 minus 46.5) with 41 (84 minus 43) degrees of freedom. The corresponding p-value is 0.02% indicating a very strong rejection of perfect risk sharing among stockholders.

On the other hand, for non-stockholders the overidentifying restriction test has a p-value slightly higher than 5% surprisingly providing much weaker evidence against perfect insurance among this group (column 1C). However, notice that all this discussion is based on estimation results without correcting for the potential selection effect. In fact, peeking ahead to Table 8 (columns 1 and 4), we see that the parameter estimates are erratic, raising suspicion about the reliability of estimation with naive GMM.

Starting with stockholders, columns 2B and 3B display the results for $h_N = 0.5$, and for the optimal bandwidth, $h_N^* = 0.24$ respectively. With bias correction, the p-value barely moves to 0.07% and is less than 0.04% at the optimal bandwidth, confirming the strong rejection for risk-sharing among stockholders. On the other hand, turning to non-stockholders (2C and 3C) something quite unexpected happens: As the bias is eliminated the weak rejection disappears and the p-value rises from 5.7% all the way up to 55%! When we further tighten the bandwidth ($h_N = 0.1$) stockholders' p-value slightly increases to 2.4% (which may partly reflect the smaller effective sample size—weights fall exponentially with h_N—and the reduced power of the test). Again, non-stockholders' p-value is unaffected (4B and 4C).

As explained earlier, it is also possible to estimate the PRS equation alone. This approach has the following advantage: a possible measurement error in wages will appear both in the error term of the MRS equation as well as in the explanatory variables of the selection equation violating assumption A1. However, the model specification test for the MRS equation (the $J-$test in 1A to 3A of Table 2) never rejects that moment condition indicating that this is not likely to be a serious

Table 2: Test of Risk Sharing with Different Bandwidth Values

	(1A)	(1B)	(1C)	(2A)	(2B)	(2C)
	Naive GMM			$h_N = 0.5$		
Group	*S+N*	*S*	*N*	*S+N*	*S*	*N*
Moment Conditions	MRS	MRS & PRS	MRS & PRS	MRS	MRS & PRS	MRS & PRS
Test Statistics						
χ^2	46.5	127.6	102.9	40.8	116.3	79.9
df	43	84	84	43	84	84
p-value (*Model*)	0.329	0.001	0.064	0.566	0.011	0.601
p-value (*PRS*)		0.0002	0.057		0.0007	0.551

Table 2 (*continued*)

	(3A)	(3B)	(3C)	(4A)	(4B)	(4C)
	$h_N^* = 0.24$			$h_N = 0.10$		
Group	*S+N*	*S*	*N*	*S+N*	*S*	*N*
Moment Conditions	MRS	MRS & PRS	MRS & PRS	MRS	MRS & PRS	MRS & PRS
Test Statistics						
χ^2	40.69	117.8	80.12	48.23	108.9	89.2
df	43	84	84	43	84	84
p-value (*Model*)	0.57	0.008	0.599	0.269	0.056	0.266
p-value (*PRS*)		0.0004	0.542		0.024	0.474

Notes: S and N denote stockholders and non-stockholders respectively. P-value (Model) is the significance level of the J-test for the joint estimation of all the moment conditions in a given column. P-value (PRS) refers to the significance level associated with the PRS moment condition df is the total degrees of freedom for all the moment conditions in a given column. The instrument set for the MRS equation includes a constant, age of head, age of spouse, household size and its lag, consumption growth lagged twice, husband's and spouse's wage growth lagged twice, and a dummy indicating stock ownership. Instrument set for PRS: include log wage change from $t-2$ to $t-1$ and from $t-1$ to t, but exclude the stockholding dummy, and use change in household size instead its levels.

Table 3: TESTS OF RISK SHARING USING THE PRS EQUATION ONLY

	(1a)	(1b)	(3a)	(3b)
	$h_N = 0.5$		$h_N^* = 0.24$	
Group	S	N	S	N
Moment Conditions	PRS	PRS	PRS	PRS
Test Statistics				
χ^2	47.68	29.54	52.07	24.47
df	29	29	29	29
p-value (PRS)	0.002	0.343	0.003	0.267

For explanations see the notes after Table 2.

problem. Moreover, this correlation is not likely to explain the asymmetry in test results, because measurement error would presumably affect both groups to similar extents. Nevertheless, to rule out measurement error completely, it is still desirable repeat the test using the PRS equation only, where head's wage does not appear.

Using (9') only PRS is still rejected for stockholders with p-values less than 0.3% (Table 3). On the other hand, for non-stockholders, the corresponding p-values are 34.3% and 26.7%, again far away from rejection.

Another advantage of this last test is that it is not affected by the identifying assumption we made about the constancy of ρ_2 and ρ_3 across the two groups. Because the PRS equation is estimated for one group at a time, that restriction is void here and the decisive rejection in this case only serves to strengthen the results.

ROBUSTNESS

These findings seem somewhat unexpected. Before discussing potential explanations for these results, let us first take a closer look and check the robustness of these results. First, endogeneity of instruments is not likely to explain these findings, because in that case the MRS equation would also be rejected when estimated alone, which was not the case. Moreover, it is not clear why invalid instruments would affect stockholders substantially while not being revealed in non-stockholders' estimation at all. Second, could this result be due to the poor finite sample properties of the GMM estimator?[22] To investigate this possibility, we reduce the degrees of freedom by eliminating lagged consumption change ($\Delta c_{n,t-2}$), and head's and spouse's lagged wage changes ($\Delta w_{1n,t-2}, \Delta w_{2n,t-2}$) from the instrument sets of both equations reducing the degree of freedom to 25. Test results reported in Table 4 confirm our previous findings. Alternatively, eliminating ($\Delta w_{1n,t}, \Delta w_{1n,t-1}$)

[22] The Monte Carlo evidence in Kyriazidou (1997, 2001) suggests that the small sample properties are quite well-behaved especially for sample sizes around what we consider in this paper. Still, we find it compelling to investigate if our results are robust to the number of instruments used.

Table 4: Further Tests of Risk Sharing Using a Small Instrument Set

	(1a)	(1b)	(1c)	(1a)	(1b)	(1c)
		$h_N = 0.5$			$h_N^* = 0.24$	
Group	—	S	N	—	S	N
Moment Conditions	MRS	MRS & PRS	MRS & PRS	MRS	MRS & PRS	MRS & PRS
Test Statistics						
χ^2	31.5	82.8	52.6	34.4	80.8	57.6
df	35	60	60	35	60	60
p-value (*Model*)	0.636	0.029	0.731	0.497	0.038	0.564
p-value (*PRS*)		0.001	0.687		0.004	0.566

The instrument set for MRS equation includes a constant, age of head, age of spouse, household size and its lag, husband's contemporaneous wage growth and its first lag. The instrument set for PRS: exclude stockholding dummy, and use change in household size instead of its level.

instead of lagged wage changes (not reported) has no appreciable effect on these results.

There are two important points that deserve further discussion. First, as noted by Hayashi et. al. (1996), tests of perfect risk sharing may not have high power against the alternative of self-insurance if the instrument set only includes lagged values of variables such as income, and exclude contemporaneous values. This is because even with incomplete markets, the permanent income hypothesis implies that lagged endogenous variables will be uncorrelated with current forecast errors. If, further, the forecast error can be written as the sum of an aggregate and an idiosyncratic component, then lagged variables will have zero correlation with the idiosyncratic component even when markets are incomplete. Hayashi et. al. could not use $\Delta w_{1n,t}$ as an instrument due to measurement error since the same variable also appears in their risk sharing equation. In contrast, our instrument set *does* include contemporaneous wage growth because $\Delta w_{1n,t}$ does not simultaneously appear in the PRS equation (which is also true in most other papers). Nevertheless, for completeness we follow Hayashi et. al and replace $\Delta w_{1n,t}$ and $\Delta w_{1n,t-1}$ with the bracketing wage changes from $t-2$ to $t+1$ and from $t-3$ to $t+1$. As Table 5 displays, PRS is still strongly rejected for stockholders, and although the p-value is somewhat lower than before for non-stockholders (=25.7% at h_N^*) it is still far away from rejection.

A second observation, originally made by Attanasio and Davis (1996), is that the tests of PRS may have higher power if one considers marginal utility growth over longer horizons than one year. This would be especially true if yearly changes in income are partly anticipated, or if they are dominated by measurement error. To implement a test with long time changes, we replace the yearly changes in the PRS equation with 6-year changes from 1985-91 and 1986-92. Also note

Table 5: Test of Risk-Sharing Using Lead Instruments

	(1a)	(1b)	(2a)	(2b)
	$h_N = 0.5$		$h_N^* = 0.24$	
Group	S	N	S	N
Moment Conditions	PRS	PRS	PRS	PRS
Test Statistics				
χ^2	57.7	40.3	51.8	36.7
df	32	32	32	32
p-value (PRS)	0.003	0.149	0.014	0.257

Notes: The instrument set includes a constant, age of head, age of spouse, change in household size, husband's log wage change from $t-3$ to $t+1$, and from $t-2$ to $t+1$.

Table 6: Test of Risk-Sharing Using Long Time Differences

	(1a)	(1b)	(2a)	(2b)
	$h_N = 3.0$		$h_N^* = 1.86$	
Group	S	N	S	N
Moment Conditions	PRS	PRS	PRS	PRS
Test Statistics				
χ^2	28.3	9.08	26.9	10.51
df	14	14	16	16
p-value (PRS)	0.012	0.825	0.019	0.720

Notes: The instrument set includes a constant, age of head, age of spouse, change in household size, consumption growth lagged twice, husband's wage growth from $t-6$ to t, and from t to $t+1$

that the optimal bandwidth has to be adjusted to account for six-year changes in wages. Table 6 displays the new results. PRS is reject at 2% significance level or lower for stockholders whereas there is still no evidence against risk sharing among non-stockholders.

After discussing the parameter estimates, we will present additional tests using a new moment restriction which essentially confirm the findings of this section. Overall, then we conclude that there is significant evidence against perfect risk-sharing among stockholders in PSID, but there is virtually no evidence against it among non-stockholders.

Before closing this section we compare our findings to existing work reviewed in the Introduction, which strongly rejected perfect insurance in the whole population. We repeat our main test of PRS for the whole population. In columns 1A and 2A of Table 7, perfect insurance is rejected at any significance level above 2%. Adding the wage equation (18) leads to even stronger rejections

Table 7: TESTS OF RISK SHARING IN THE WHOLE POPULATION

	(1a)	(1b)	(1a)	(1b)
	$h_N = 0.50$		$h_N^* = 0.24$	
Group	S+N	S+N	S+N	S+N
Moment Conditions	MRS & PRS	MRS & PRS & WAGE	MRS &PRS	MRS & PRS & WAGE
Test Statistics				
χ^2	86.35	137.67	86.07	135.4
df	71	111	71	111
p-value (*Model*)	0.104	0.043	0.017	0.057
p-value (*PRS*)	0.0195	0.001	0.020	0.003

Notes: The instrument set is the same one as in Table 2. See the notes after Table 2 for details.

(columns 1B and 2B).[23] In light of this finding, it seems that the failure of PRS in the whole population noted in the literature is likely to be due to the failure of risk-sharing not among the poor, but among the wealthy.

7.2 Parameter estimates

All the structural parameters of the model can be identified by jointly estimating: (1) the MRS equation, and (2) the PRS condition for either stockholders or non-stockholders. The wage equation (18) is added as a third moment condition to obtain more precise estimates. In Table 8, the first three columns report the estimates obtained using stockholders' PRS conditions using the naive GMM estimator (column 1), the kernel-weighted GMM with $h_N = 0.5$ and the optimal bandwidth, $h_N^* = 0.24$ (columns 2 and 3). Similarly, the last three columns report the corresponding estimates obtained by using non-stockholders' PRS equation.

As before, we begin by analyzing the effect of bias correction. First, when the model is rejected (first four columns), as can be expected, many parameters have the wrong sign: for example, in columns 2 to 4, $\rho_2, \rho_3 \gg 1$ implying that preferences are convex in female leisure time. Also in these same columns, a new child is apparently not a welcome guest: the household size coefficients B_0^3 and B_1^3 are often negative, meaning that an increase in household size, which is mainly due to a new child in our sample of married couples, decreases utility everything else held constant. Also, in the first four columns standard errors are often very large. In the rest of this section we will focus our discussion on the last two columns where the underlying model is not rejected (i.e.,

[23]Note that because switchers ($d_{nt} \neq d_{n,t-1}$) have been eliminated from the sample previously, we still use the kernel-weighted estimator.

Table 8: STRUCTURAL PARAMETERS FROM THE JOINT ESTIMATION OF MRS, PRS AND WAGE EQUATIONS

	$U^i = \delta_0(\mathbf{Z})C^{\rho_0^i}L_2^{\rho_2} + \delta_1(\mathbf{Z})L_1^{\rho_1^i}L_2^{\rho_3}$					
	(1)	(2)	(3)	(4)	(5)	(6)
Group	STOCKHOLDERS			NON-STOCKHOLDERS		
Bandwidth	*nGMM*	$h_N = 0.5$	$h_N^* = 0.24$	*nGMM*	$h_N = 0.5$	$h_N^* = 0.24$
Rejected at 5%?	**Yes**	**Yes**	**Yes**	**Yes**	**No**	**No**
Curvature Parameters						
ρ_0^{no}	0.69	−0.58	0.42	−0.55	−0.24	−0.18
	(0.56)	(4.78)	(0.71)	(7.23)	(0.39)	(0.48)
ρ_1^{no}	−12.9	−13.3	−6.07	−8.02	−5.53	−5.46
	(2.98)	(32.26)	(3.21)	(30.61)	(1.15)	(1.53)
ρ_2	−23.88	95.25	12.6	156.44	−32.9	−35.7
	(16.0)	(132.3)	(34.8)	(436.6)	(32.4)	(22.8)
ρ_3	−57.2	61.92	−5.78	123.1	−45.7	−50.1
	(14.8)	(52.2)	(33.4)	(539.4)	(33.9)	(26.2)
$a_0 \left(= \rho_0^{st} - \rho_0^{no}\right)$	4.21	3.69	0.95	3.71	1.65	1.84
	(1.05)	(10.2)	(1.29)	(14.6)	(0.72)	(1.02)
$a_1 \left(= \rho_1^{st} - \rho_1^{no}\right)$	−4.51	−9.8	−0.21	−7.12	−6.83	−10.1
	(3.86)	(29.8)	(6.52)	(35.3)	(3.30)	(4.12)
Demographic Effects						
B_0^1 (age-squared)	1.12	−6.4	0.30	3.34	−1.29	−1.92
	(3.72)	(13.5)	(2.97)	(12.49)	(3.01)	(3.96)
B_0^2 (age-cubed)	2.49	13.2	1.15	8.72	−3.48	−5.21
	(1.18)	(9.81)	(1.15)	(7.68)	(1.32)	(1.73)
B_0^3 (family size)	−0.46	2.64	−0.31	−2.60	0.93	1.39
	(0.99)	(9.29)	(1.02)	(8.88)	(1.39)	(0.92)
B_1^1 (age-squared)	0.71	−6.76	0.17	2.92	−1.48	−2.17
	(3.72)	(13.2)	(2.99)	(12.88)	(2.82)	(3.76)
B_1^2 (age-cubed)	2.65	−13.11	1.24	8.91	−3.41	−5.16
	(3.72)	(13.5)	(2.98)	(12.45)	(2.53)	(3.35)
B_1^3 (family size)	1.10	4.14	0.45	−1.05	1.53	2.04
	(3.66)	(14.77)	(3.01)	(13.12)	(3.01)	(3.58)
Test Statistics						
χ^2 (*Model*)	170.9	158.3	159.8	145.1	128.3	122.3
df (Model)	124	124	124	124	124	124
p-value (*Model*)	0.003	0.021	0.016	0.069	0.371	0.536
p-value (*PRS*)	0.0001	0.0006	0.0004	0.047	0.462	0.429

Notes: nGMM denotes the naive GMM estimator which does not correct for selection bias ($h_N = \infty$). The structural parameters are exactly identified and the standard error for parameter estimates are in paranthesis.

non-stockholders PRS equation is imposed). In this latter case, the estimates are economically more sensible and are estimated much more precisely. This difference also stresses the role of bias correction in bringing the model closer to data.

The curvature of head's leisure in non-stockholders' utility function, ρ_1^{no}, is quite precisely estimated around -5 to -6 implying that utility is concave and increasing in leisure as predicted by theory (recall that the second subutility is weighted by $1/\rho_1^{no}$ so that it is increasing in leisure). Stockholders' curvature parameter, $\rho_1^{no}+a_1$, is significantly more negative: around -13 to -15. Moreover, from the reduced form coefficients, the null restriction $a_1/(\rho_3-\rho_2)=0$ has a $t-$distribution with 1 *df* and its value is 4.2 strongly rejecting no heterogeneity across the two groups.

An economically more meaningful measure can be obtained from these parameters: the elasticity of male labor supply with respect to wages (holding the marginal utility of wealth constant) is $L_{1nt}\left((1-L_{1nt})(1-\rho_1)\right)^{-1}$, which can easily be derived from the first order condition for labor choice, equation (6). Given that the average time spent at work is approximately 2200 hours per year for both groups in our sample, assuming 16 hours of discretionary time per day, we get $L_{1nt}=0.37$. Then the implied elasticities are 0.21 and 0.10 for stockholders and non-stockholders respectively. This value is well within the range of numbers reported in the literature (see Pencavel (1986) for a survey). Furthermore, given that stockholders are on average better educated (Table 1), these elasticity figures also support some previous empirical evidence that better educated men have a lower labor supply elasticity. As for the curvature parameters for female leisure, ρ_2 and ρ_3, the point estimates are always negative which implies that preferences are again concave and increasing as predicted by theory.[24]

The curvature coefficient for consumption in non-stockholders' utility function, ρ_0^{no}, is estimated quite precisely to be around zero in all columns implying logarithmic preferences in consumption. The difference between the curvature coefficients of the two groups $\left(a_0=\rho_0^{st}-\rho_0^{no}\approx 1.6 \text{ to } 1.8\right)$ is statistically significant. Using the reduced form coefficients from column 5, the null hypothesis $a_0/(\rho_3-\rho_2)=0$ has a $t-$distribution with 1 *df* and its value is 3.3 rejecting the null at 1% significance level or lower.

A puzzling observation is that the point estimates of stockholders' curvature parameter are greater than 1.0, which, if true, implies that their preferences are convex in consumption. This convexity has also been noted by Altug and Miller (1990) although it is not statistically significant either here or in that paper. However, given that perfect risk sharing is rejected among stockholders,

[24] One potential caveat about the preference heterogeneity revealed in ρ_0 and ρ_1 is that it may in fact be coming from heterogeneity in ρ_2 and ρ_3 across the two groups which we assumed away. This possibility can be easily checked by estimating the MRS condition *separately on each subsample* at the expense of less precise estimates due to smaller sample sizes. This estimation yields quite similar estimates for both groups ($\rho_0^{no}=-0.15$, $\rho_0^{st}=3.2$, $\rho_1^{no}=-6.7$, $\rho_1^{st}=-19.3$) suggesting that the heterogeneity uncovered is genuine.

a plausible alternative model is that stockholders' data are efficient allocations in an economy with *private information* and *long-term contracts.* Private information is likely to arise in business ownership/managerial roles which are concentrated among (the wealthy) stockholders. It is well-known that in this case, instead of the perfect risk-sharing condition, an optimality condition similar to the consumption Euler equation holds, but in the place of consumption growth, its reciprocal, $\left(\frac{c_t}{c_{t+1}}\right)$, appears (c.f., Rogerson (1985); and Ligon (1998)). Consequently, the estimated exponent has the same magnitude as the RRA coefficient but with the inverse sign, which is what we seem to have found. Hence, this curious finding seems point to the importance of private information for stockholders. We discuss this point further in the next section.

Nevertheless, it is still fair to argue that if stockholders' preferences are indeed convex in consumption and there are no asymmetric information problems, then their first order condition with respect to consumption will no longer determine optimal choices invalidating the tests of risk-sharing for that group. Fortunately, there is another equation, namely the optimality condition for male leisure choice (6) that can also be used to test risk-sharing. To investigate the severity of this problem, in the next subsection we conduct further tests of PRS implied by this equation.

Finally, the estimated coefficients of household characteristics also seem sensible. The structural coefficient of H_{nt} is positive which means that an increase in household size (which is mostly due to a new child, since our sample contains only married couples) increases both subutilities. Also, from the reduced form coefficients in the marginal utility growth equation, a new child increases female leisure time by 5% in the first year, holding other variables constant. Considering that the average female in our sample works for approximately 1500 hours a year (conditional on working), this implies 220 hours, or equivalently, seven weeks of work hours reduction in her labor supply over the first year, which probably seems reasonable.[25]

7.3 Further tests of risk-sharing

To investigate if the possible non-concavity of utility function might have affected our results, we conduct additional tests. Perfect risk-sharing also imposes structure on the cross-section of marginal utility of leisure growth. After taking logs, differencing and rearranging equation (6), we get

$$-\rho_2 \Delta l_{2nt} = -\Delta w_{1nt} - \Delta \ln(\lambda_t) + \mathbf{B}_0 \Delta x_{nt} + (\rho_1 + a_1 - 1)\Delta l_{1nt} + \Delta \nu_{4nt} \qquad (19)$$

where the error term, $\Delta\nu_{4nt}$, includes the same components as $\Delta\nu_{1nt}$. As can be seen in Table 9, the results are very similar: risk-sharing is rejected ($p-value = 0.7\%$) for stockholders, but not for non-stockholders ($p - value = 64.8\%$). Adding the MRS equation as before yields an even lower

[25] Since the sample also includes women who do not work during the estimation period, their leisure hours do not decrease after the birth of a baby. Hence, this figure (7 weeks) is likely to be a lower bound for working women.

Table 9: TESTS OF RISK SHARING USING THE MARGINAL UTILITY OF LEISURE EQUATION

	$h_N^* = 0.24$				
Group	—	STOCKHOLDERS		NONSTOCKHOLDERS	
Moment Conditions	MRS	PRS	MRS & PRS	PRS	MRS & PRS
Test Statistics					
χ^2	40.69	49.23	97.57	24.62	62.03
df	43	28	71	28	71
p-value (*Model*)	0.571	0.008	0.019	0.648	0.767
p-value (*PRS*)		0.007	0.001	0.648	0.815

Notes: The instrument set is the same one as in Table 2 but excludes contemporaneous wage change since it appears in (6). See the notes after Table 2 for other details.

p-value of 0.1% for stockholders and again no rejection for non-stockholders.

The fact that the test results are pretty much unchanged supports our argument that we should probably not read too much into the point estimates of $(\rho_0 + a_0)$ which are not significantly greater than unity, and that the true value is probably smaller than one. Hence the first order condition for consumption is also very likely to be valid.

8 Discussion and Conclusion

In this paper we found strong evidence against risk-sharing among wealthy stockholders, but found virtually no evidence against it among non-stockholders. This result is robust to a number of changes made, such as including future wages into the instrument set and testing from long time differences of moment conditions, which are emphasized in the previous literature. Moreover, risk-sharing in the whole population is strongly rejected in this paper consistent with existing literature. Overall, these findings suggest that the failure of PRS in the whole population is likely to be due to the failure of the wealthy to insure the additional risks they face.

These results appear surprising until one takes into account the different risk situations faced by each group. As noted earlier, households in the top 20 percent of the wealth distribution own more than 90 percent of non-housing wealth—including private capital—and almost all the corporate equity outstanding. Thus, these households are the main investors/business owners in the economy and are thus exposed to (idiosyncratic) production risk.[26]

The large literature on corporate ownership structure as well as the more recent work on private business ownership find agency costs (moral hazard) to be empirically important sources of market

[26] According to the 1995 Survey of Consumer Finances the fraction of business owners (with a business value above $10,000) among non-stockholders is only 3 percent.

incompleteness.[27] Thus, in an environment with private information, active owners/managers of business assets will be exposed to idiosyncratic (undiversified) risk for incentive reasons. This is true not only for small entrepreneurs but also for the corporate ownership structure both of which directly affect wealthy households.[28] In other words, although the wealthy have more insurance opportunities, their incomes are significantly harder to insure as well. The wage measure used in this paper includes labor portion of business income and can thus capture the correlation between consumption growth and shocks to business income.

On the other hand, the large majority of non-stockholders are simply workers whose main source of income are wages. A number of implicit or explicit sources of insurance are already built into this income: long-term contracts, welfare programs, minimum wage laws, and unemployment insurance which are usually extended during downturns, are all designed to insure workers from fluctuations in business conditions. In addition, a number of informal risk sharing mechanisms (e.g., inter-vivos transfers, charitable donations, and borrowing and lending) further eliminate the risks faced by most households. Note that many of these insurance opportunities are not likely to be effective in insuring losses experienced by business owners. Overall these results underscore the importance of risks faced by the wealthy as important sources of market incompleteness.

To deal with the selection problem arising from the endogenous stock market participation choice, we implemented a powerful new semiparametric GMM estimator proposed by Kyriazidou (2001) for the first time. One conclusion that we draw is that self-selection into the stock market seriously biases the results if not corrected for. For example, without correction the parameter estimates are erratic, and often have wrong signs. In this sense, this analysis bears witness to the usefulness of this estimator. Also, the weak rejection of PRS for non-stockholders is overturned when the model is estimated correcting for the bias, although PRS is rejected in all cases for stockholders.

From a substantive viewpoint, these results suggest that stockholders and non-stockholders seem to face different risk situations and have tools of varying effectiveness to insure against these shocks. As a result, the importance of market incompleteness and idiosyncratic shocks are different for the wealthy and the average household.

[27] There is a large literature on the corporate ownership structure. The classic reference is Jensen and Meckling (1976); see also Himmelberg, Hubbard, Love (2002) and the references therein. On entrepreneurial income risk, see Bitler, Moskowitz and Vissing-Jorgensen (2002)

[28] Although it is beyond the scope of this paper, in principle it is possible to identify the role of entrepreneurial risk for stockholders. In particular, looking at stockholders only, one can test the PRS separately for those who own private businesses and those who do not. If risk-sharing is rejected for the former group but not for the latter, then this would suggest entrepreneurial risk as an important source of market incompleteness. However, entrepreneurship introduces another selection equation which must be dealt with as well. So, we leave this question for future research.

A Appendix: Derivation of the Risk-Sharing Condition

It is easier to derive the risk-sharing condition from the sequential formulation of the decision problem rather than the recursive formulation stated in the text. We use the same notation here as in Section 2. For clarity, we specialize to the case where households only derive utility from consumption; as will become clear, the derivation extends to the more general u considered in the text straightforwardly.

Recall that s_t denotes a particular state at time t, and $s^t = (s_1, ..., s_t)$ is the history of states up to and including time t. As usual the choice objects for time t are $\mathcal{F}_t$−measurable random variables. Let $\pi(s^t)$ denote the probability of history s^t being realized conditional on time zero information. After each history, each household makes a stock market participation choice, $d(s^t)$, choose consumption, $C(s^t)$, and a portfolio choice vector, $\mathbf{k}(s^t)$ if $d(s^t) = 1$, or $k_0(s^t)$ if $d(s^t) = 0$. We can think of an agent's choice as a two step procedure. In the first step, assign a participation choice for every possible history, and for this given sequence solve the optimal consumption and portfolio choice. Repeat this for all possible sequences of participation choices. In the second step pick the combination of participation choice sequence which yields the highest lifetime utility. The maximum exists because of finite lifetimes. Denote this optimal decision by $d^*(s^t)$. Given this optimal participation choice, after any history $s^{t'}$, and for $t \geq t'$, households face the following problem:

$$\max_{C(s^t),\mathbf{k}(s^t),k_0(s^t)} \left[\sum_{s^t : t \geq t'} \pi\left(s^t | s^{t'}\right) u\left(C\left(s^t\right)\right) \right]$$

s.t

$$\begin{aligned} C\left(s^t\right) + \sum_{j=1}^{S} k_j\left(s^t\right) &= \omega\left(s^t\right) \\ \omega\left(s^{t+1}\right) &= \sum_{j=1}^{S} k_j\left(s^t\right)\left(1 + R_j\left(s^{t+1}\right)\right) \end{aligned}$$

for histories such that $d^*(s^t) = 1$, and

$$\begin{aligned} C\left(s^t\right) + k_0\left(s^t\right) &= \omega\left(s^t\right) \\ \omega\left(s^{t+1}\right) &= k_0\left(s^t\right)\left(1 + R_0\right) \end{aligned}$$

for histories such that $d^*(s^t) = 0$.

Consider a current stockholder, $d^*\left(s^{t'}\right) = 1$, and without loss of generality, assume that $d^*\left(s^{t'+1}\right) = 1$, for the first s^* possible states in the next period, and $d^*\left(s^{t'+1}\right) = 1$ for the remaining $S - s^*$ states. By substituting the budget constraints at t' and $t'+1$ into the objective function we get:

$$\begin{aligned} &\max_{C(s^t),\mathbf{k}(s^t),k_0(s^t)} u\left(\omega\left(s^t\right) - \sum_{j=1}^{S} k_j\left(s^t\right)\right) + \sum_{i=1}^{s^*} \pi\left(s_i^{t'+1}|s^{t'}\right) u\left(\sum_{j=1}^{S} k_j\left(s^t\right)\left(1+R_j\left(s_i^{t+1}\right)\right) - \sum_{j=1}^{S} k_j\left(s_i^{t'+1}\right)\right) \\ &+ \sum_{i=s^*+1}^{S} \pi\left(s_i^{t'+1}|s^{t'}\right) u\left(\sum_{j=1}^{S} k_j\left(s^t\right)\left(1+R_j\left(s_i^{t+1}\right)\right) - k_0\left(s_i^{t'+1}\right)\right) + \sum_{s^t : t > t'+1} \pi\left(s^t|s^{t'+1}\right) u\left(C\left(s^t\right)\right) \end{aligned}$$

The second term is the expected utility in the next period over states where stockholding is optimal. Similarly, the third term is the expected utility over states where non-stockholding is optimal. Note that the only difference between these two terms is the investment decision in the next period: $\sum_{j=1}^{S} k_j\left(s_i^{t'+1}\right)$ versus $k_0\left(s_i^{t'+1}\right)$. Finally, the last term captures remaining lifetime utility after period $t'+1$.

The first order conditions for portfolio choice at time t' for asset j is given by differentiation with respect to $k_j(s^t)$:

$$0 = -u'\left(C\left(s^{t'}\right)\right) + \sum_{i=1}^{S} \pi\left(s_i^{t'+1}|s^{t'}\right) u'\left(C\left(s_i^{t'+1}\right)\right)\left(1+R_j\left(s_i^{t+1}\right)\right), \quad \text{for } j = 1, ..., S$$

Since the different terms $\left(\sum_{j=1}^{S} k_j\left(s_i^{t'+1}\right) \text{ versus } k_0\left(s_i^{t'+1}\right)\right)$ do not appear in the derivative of u, we combined the two summations and obtain the standard Euler equation for consumption allocation. Also, note that this same equation holds regardless of the value of s^* or the ordering of states, so that households who may find it optimal to participate in different states tomorrow will have the same equation. Rearrange to obtain:

$$\sum_{i=1}^{S} \pi \left(s_i^{t'+1} | s^{t'}\right) \frac{u'\left(C\left(s_i^{t'+1}\right)\right)}{u'\left(C\left(s^{t'}\right)\right)} \left(1 + R_j\left(s_i^{t+1}\right)\right) = 0, \qquad \text{for } j = 1, ..., S$$

Just like in the standard case with complete markets in all periods, the S equations above give a unique solution for the S marginal utility growths in terms of aggregate variables. Thus, marginal utility growth does not depend on individual variables, implying perfect risk-sharing:

$$\frac{u'\left(C\left(s^{t'+1}\right)\right)}{u'\left(C\left(s^{t'}\right)\right)} = \frac{\lambda_{t+1}}{\lambda_t}$$

where λ denotes this aggregate shock.

It is clear that this derivation does not depend on u being only a function of consumption. Moreover, following the same approach, it is easy to show that a similar condition holds for marginal utility of leisure growth (given in equation (5)).

B Appendix: The Data

We use the PSID data set on U.S. households. Starting with the "family files" from 1982 to 1993 waves, we use the following sample selection criteria to select our main sample. Specifically, we include household-years in t and $(t-1)$ in estimation if the head of the family:

(i) is in the study for at least four consecutive years $(t-3, t-2, t-1, t)$ including 1984 or 1989,

(ii) is married to the same spouse at least in the last two years $(t-1, t)$ of the same period,

(iii) has a positive labor income at least in the last two years $(t-1, t)$ of the same period.

These criteria produced a sample of 2350 households who were in the study between 1984 and 1993, not necessarily for all years. Further, we eliminated households who did not satisfy data reliability controls on some key variables as follows:

We have eliminated a household-year if:

(iv) annual family food consumption expenditure was less than $150,

(v) head's education variable was missing for the last two years $(t-1, t)$ of this period,

(vi) if head's or spouse's reported annual labor hours exceeded 4860 hours.

Criteria (i) and (iii) above are used in most analysis of labor or consumption data, (see Altonji, 1986) to eliminate irregular observations.

(vii) Finally if a household changed its stockholding status from 1984 to 1989, we eliminate that observation from estimation between these two dates..

Apart from these we encountered a few cases where head or spouse had positive annual labor hours but zero annual labor income, or vice versa. These observations were also filtered out. In PSID most variables have top coding. We also eliminate a household-year if the upper bounds for consumption is binding.

Another important concern is coding errors. Although, it is not possible to identify all of them, there is one type which is not very hard to detect, and which can also seriously affect the tests of risk-sharing: sometimes there is an omission or an addition of an extra digit during coding. This results in a large sudden jump or drop in the time-series of that variable. Of course if the variable has very large variance, observed fluctuations may also be genuine, and eliminating them may reduce the power of our tests. Thus, we first isolated observations on consumption and head's and spouse's wages which violated the following bound: $E(X_t) - 2 * std(X_t) \leq X_t \leq 2 * E(X_t) + 2 * std(X_t)$. This is clearly a generous bound. There were a total of 46 observations which violated this bound for at least one of the three variables. Upon closer inspection of the time-series of these variables, we eliminated 41 observations which had small standard deviations and the outlier was very close to 10 times (or 0.1 times) the sample average.

These criteria produced the following number of observations in each year:

Number of Observations for Moment Conditions of Time Difference: Year (t) - Year ($t+1$)							
Moment Condition	$83-84$	$84-85$	$85-86$	$89-90$	$90-91$	$91-92$	*TOTAL*
# of observations	1292	1289	1302	1761	1709	1588	8941

Wages: The average hourly labor earnings (wages) of head and spouse reported in PSID and adopted in this paper are calculated from the sum of the following types of income and total annual hours:

V19127 Labor Part of Farm Income
V19128 Labor Part of Business Income
V19129 Salary Income
V19131 Bonuses, Overtime, Commissions
V19132 Income from Professional Practice or Trade
V19133 Labor Part of Market Gardening Income
V19134 Labor Part of Roomers and Boarders Income.

Stockholding: The definition of stockholding adopted in this paper includes ownership of shares of stock in publicly held corporations, mutual funds, investment trusts, including stocks in IRA's. This definition corresponds to PSID variables V10912 for 1984 and V17325 for 1989. All households who indicate they do not own any of these assets are considered non-stockholders that year.

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