Real Exchange Rates with Limited Asset Market Participation: An Analysis with Micro Data

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Abstract

Standard models of open-economy macroeconomics predict a strong correlation between real exchange rates and relative consumption across countries. The predicted relationship fails overwhelmingly when confronted with data. A promising way to break the link between real exchange rates and aggregate consumption is to account for heterogeneity across households due to differences in asset market participation status. This paper studies the equilibrium condition that arises when only a fraction of households participate in asset markets in each period. In such an environment, real exchange rates are related to the consumption of the households that participate in asset markets instead of the aggregate consumption. We test this condition for real exchange rates between nineteen administrative regions of Italy using micro data on household consumption and financial assets from the Survey of Household Wealth and Income collected by the Bank of Italy. Real exchange rates are constructed using regional price indices collected by the Italian National Institute of Statistics. The nonlinear GMM estimation of moment conditions provides evidence in favor of the model's implications. In particular, overidentification restrictions of the model are not rejected at conventional levels of significance. The evidence with respect to parameter estimates is more mixed yielding positive and statistically significant risk aversion parameters for one third of the regions we consider.

Key Words: Limited asset market participation, discrete-choice models
JEL Classification: C25, G11

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1 Introduction

Standard models of open-economy macroeconomics predict the correlation between real exchange rates and relative consumption across countries to be close to one. The implication that a country that experiences a real appreciation should have a corresponding decrease in its relative consumption is rejected overwhelmingly in the data. This so-called consumption-exchange rate anomaly is documented by Backus and Smith (1993) using time-series data for eight OECD countries.

Several extensions to the standard model have been proposed to account for this delink between real exchange rates and relative consumption across countries. Suggested solutions include models that introduce various frictions in goods markets as well as asset markets. For example, Backus and Smith (1993) study the possibility that non-traded goods might help bridge the gap between predictions of the theory and the empirical evidence. They conclude that even with the presence of non-traded goods the correlation between relative price movements and relative consumption movements across countries remains low. Chari, Kehoe and McGrattan (2002) abstract from non-traded goods and rely on price stickiness for traded goods to generate the observed characteristics of real exchange rates and relative consumption. Such a market structure yields fluctuations in real exchange rates that are similar to those observed in the data. However, their model predicts a high correlation between real exchange rates and relative consumption in sharp contrast to the empirical evidence. Both of these studies list incomplete asset markets as a promising avenue to resolve this puzzle. In fact, Chari et al. (2002) introduce incomplete markets by restricting assets traded across countries to a single uncontingent nominal bond. This change fails to dispel the anomaly by leaving the correlation between real exchange rates and relative consumption almost the same as in their benchmark model. Corsetti, Dedola, and Leduc (2003) combine a similar friction in asset markets with a different friction in goods market, distribution services intensive in local inputs. Under this market structure, the predicted correlation between real exchange rates and relative consumption turns out to be negative conditional on a productivity increase in tradable goods.

All these studies shed light on different aspects of the consumption-exchange rate puzzle formulated by the following equation:

$$\frac{e_t P_{jt}}{P_{it}} = \frac{U'(c_{jt})}{U'(c_{it})}, \tag{1}$$

where e is the nominal exchange rate, and P_i and P_j are the price levels in countries i and j respectively. $U'(c_i)$ denotes the marginal utility with respect to consumption for country i and $U'(c_j)$ is the marginal utility with respect to consumption for country j. However, a satisfactory explanation of the puzzle consistent with the data is still an open question.

An alternative and promising way of breaking the link between real exchange rates and relative consumption is to account for limited asset market participation. Acknowledging that a substantial fraction of households do not participate in asset markets is aimed at replacing the representative household assumption as one of the building blocks of standard models that give way to the above equation. There are several advantages to this approach. First of all, limited asset market participation is a characteristic of even some of the most developed financial systems around the globe. Figures based on various financial assets from different economies help establish this phenomenon as a prominent feature of asset market structure. Haliassos and Bertaut (1995) report that data from the 1983 Survey of Consumer Finances (SCF) reveal the proportion of households in United States that hold stocks either directly or through mutual funds to be well below 50 percent except for the top quintile of income. Using the Panel Study of Income Dynamics (PSID), Mankiw and Zeldes (1991) also document that only a small proportion of households holds stocks. In a more recent study, Vissing-Jørgensen (2002) uses the Consumer Expenditure Survey (CEX) where she classifies 21.75 percent of U.S. households as stockholders. A more comprehensive indication of limited asset market participation is provided by Mulligan and Sala-i-Martin (2000). They use data from the 1989 Survey of Consumer Finances to document that 59 percent of U.S. households do not hold any nonmonetary financial assets. These interest-bearing financial assets include money-market accounts, CDs, other bonds, mutual fund shares, and equities. Limited asset market participation is evident in other developed economies as well. For example, Attanasio, Banks and Tanner (2002) state that more than 75 percent of U.K. households do not own shares directly while a level of share ownership greater than 20 is a recent phenomenon.

It is therefore no surprise that there is an ever growing empirical literature on the implications of limited asset market participation for various aspects of the economy. A particularly relevant strand of the literature emphasizes the differences in consumption processes between asset holders and non-asset holders. For example, Attanasio, Banks and Tanner (2002) investigate whether accounting for limited participation can explain the lack of correspondence between the sample moments of the intertemporal marginal rate of substitution and asset returns in the U.K. data. They conclude that the predictions of the consumption capital asset pricing model cannot be rejected for share holders, in contrast to the failure of the model when estimated on data for all households. Vissing-Jørgensen (2002) provides empirical evidence from the U.S. household data that accounting for limited asset market participation is important for estimating the elasticity of intertemporal substitution, one of the central determinants of households' intertemporal consumption choice. Güvenen (2003) offers additional support by showing that contradicting results on the value of elasticity of intertemporal substitution is partially due to ignoring household heterogeneity with respect to asset market participation status.

In this paper, we undertake an empirical investigation of whether and to what extent limited asset market participation can help account for the lack of correlation between real exchange rates and relative consumption. Our empirical tests are based on a theoretical framework by Alvarez, Atkeson and Kehoe (2002). They derive the equilibrium condition that arises when only a fraction of households participate in asset markets due to a friction introduced through a fixed cost of participation. This equilibrium condition differs from the one that arises in standard models in that the real exchange rate is not linked to a ratio of aggregate consumption across countries. Instead, it is the consumption of households that participate in asset markets that is related to the real exchange rate. More specifically, the relationship is given by the following equation:

$$\frac{e_t P_{jt}}{P_{it}} = \frac{U'(c_{Ajt})}{U'(c_{Ait})} \tag{2}$$

where e is the nominal exchange rate, and P_i and P_j are the price levels in countries i and j respectively. $U'(c_{Ai})$ denotes the marginal utility with respect to consumption for the asset market participant household in country i and similarly in country j.

Obviously testing this relationship requires the use of micro data on household consumption as well as asset holdings. Our empirical work starts off with identification of asset holders from micro household data. Estimation of a probit model reveals that factors such as age, education, geographical area of residence as well as time (trend) contribute significantly to household asset market participation decision.

Next, we undertake an empirical investigation of the relationship that relates bilateral real exchange rates to relative consumption of the asset market participants across economies. The nonlinear GMM estimation of moment conditions based on equation (2) provides evidence in favor of the model's implications. In particular, overidentification restrictions of the model are not rejected at 5% level of significance. The evidence with respect to the estimates of the risk version parameter α is more mixed yielding positive and statistically significant parameters for one third of the regions we consider.

The paper is organized as follows. Section 2 details the economic environment. Section 3 describes the data used in empirical work. The empirical specification is discussed in Section 4. Section 5 explains the identification of asset holders from the data. Results from the GMM estimation of our empirical model are summarized in Section 6. Section 7 concludes with plans for future work.

2 Economic Environment

In order to highlight the mechanisms which introduce limited asset market participation into an otherwise standard cash-in-advance framework, we start with the one-country version of Alvarez, Atkeson, and Kehoe (2002) framework. The economy is inhabited by a continuum of households of measure 1, and a government and lasts for an infinite number of time periods $t = 0, 1, 2, \ldots,$. In each period households receive an idiosyncratic endowment good y_t which is distributed according to the probability density function $f(y_t)$. The probability density function over a typical history of individual endowment shocks $y^t = (y_0, \ldots, y_t)$ is then given by $f(y^t) = f(y_0)f(y_1)\ldots f(y_t)$. As in a standard cash-in-advance model like that of Lucas (1984), trade takes place in separate goods and asset markets. In the asset market households exchange cash and bonds that promise delivery of cash in the next period. In the goods market, households use cash to buy goods subject to a cash-in-advance constraint, and sell their endowments of goods for cash. Households enter each period $t \geq 1$ with cash holdings of $P_{-1}y_{-1}$, obtained in the previous period from selling their idiosyncratic endowment y_{-1} at the price level P_{-1}. The real value of the cash holdings is determined prior to any trade when government conducts an open market operation in the asset market setting money growth μ and current price level P. The probability density over history of money growth shocks $\mu^t = (\mu_1, \ldots, \mu_t)$

is given by $g(\mu^t) = g(\mu_1)g(\mu_2)\ldots g(\mu_t)$. Following the realization of money growth and price level, each household splits into a worker and a shopper. The worker takes the current idiosyncratic endowment y and sells it for cash Py in the goods market and then rejoins the shopper at the end of the period. The shopper takes the household's cash from previous period $P_{-1}y_{-1}$ with a real value of $m = P_{-1}y_{-1}/P$ and shops for goods. The shopper also decides whether or not to pay a fixed cost of γ to transfer cash Px with real value of x between goods and asset markets. If the fixed cost is paid, then the cash-in-advance constraint is $c = m + x$ where c is consumption, and otherwise the constraint is $c = m$. Due to the binding cash-in-advance constraint households do not carry cash in the goods market from one period to the next. Instead, they save by investing in interest-bearing assets.

In the asset market, households trade a complete set of one-period bonds with payoffs contingent on both the aggregate money growth and the individual endowment shock. They enter each period t with bond holdings $B(\mu^t, y^{t-1})$ that pay \$1 contingent on relevant events. The cash proceeds from these bonds can be used to purchase new bonds $B(\mu^t, \mu_{t+1}, y^{t-1}, y_t)$ at a price given by $q(\mu^t, \mu_{t+1}, y_t)$. The cash can also be transferred to the goods market conditional on paying the fixed cost γ.

Under this market structure households' optimization problem is to choose consumption c and cash transfer x to

$$\max \sum_{t=0}^{\infty} \int \int \beta^t U'(c(\mu^t, y^{t-1}))g(\mu^t)f(y^{t-1})d\mu^t dy^{t-1} \tag{3}$$

subject to the cash-in-advance constraint

$$c(\mu^t, y^{t-1}) = m(\mu^t, y^{t-1}) + x(\mu^t, y^{t-1})z(\mu^t, y^{t-1}) \tag{4}$$

and the sequence of budget constraints for $t \geq 1$

$$
\begin{aligned}
B(\mu^t, y^{t-1}) \;\; = \;\; & \int_{\mu_{t+1}} \int_{y^t} q(\mu^t, \mu_{t+1}) B(\mu^t, \mu_{t+1}, y^{t-1}, y_t) f(y_t) dy_t d\mu_{t+1} \\
& + P(\mu^t)[x(\mu^t, y^{t-1}) + \gamma] z(\mu^t, y^{t-1}) \qquad (5)
\end{aligned}
$$

where z is an indicator function that takes on a value of 1 if the household decides to pay the fixed cost γ and transfer cash with a real value of x between asset and goods market and 0 otherwise.

In equilibrium, only a fraction of households pay the fixed cost and participate in asset markets setting their indicator function z to one. Resulting segmentation of the asset market has important consequences for determination of asset prices in this economy. To show this, the sequence of budget constraints given by equation (5) can be used to substitute out for household's bond holdings and to replace these constraint with a single period 0 budget constraint given by

$$
\sum_{t=0}^{\infty} \int Q(\mu^t) \int_{y^{t-1}} \{P(\mu^t)[x(\mu^t, y^{t-1}) + \gamma] z(\mu^t, y^{t-1})\} f(y^{t-1}) dy^{t-1} d\mu^t \;\; \leq \;\; \bar{B} \qquad (6)
$$

where $Q(\mu^t)$ denotes the price in dollars in the asset market in period 0 for a dollar delivered in the asset market in period t in state μ^t. $\bar{B}$ denotes the government debt held by the household in period zero which is a claim on $\bar{B}$ dollars in the asset market in period 0.

Then the household optimization problem is redefined as choosing consumption c and cash transfer x to maximize utility subject to the cash-in-advance constraint given by equation (4) and the period 0 budget constraint given by equation (6). With $\nu(\mu^t, y^{t-1})$ and λ as the Lagrangian multipliers on constraints (4) and (6) respectively, the two first-order-conditions for this optimization problem are given by

$$\beta^t U'(c(\mu^t, y^{t-1}))g(\mu^t)f(y^{t-1}) = \nu(\mu^t, y^{t-1}) \tag{7}$$

and

$$\lambda Q(\mu^t)P(\mu^t)z(\mu^t, y^{t-1})f(y^{t-1}) = \nu(\mu^t, y^{t-1})z(\mu^t, y^{t-1}). \tag{8}$$

It follows that when $z(\mu^t, y^{t-1}) = 1$, period 0 nominal asset prices are determined by the first-order-conditions for households who participate in asset markets by paying the fixed cost as follows:

$$\beta^t U'(c(\mu^t, y^{t-1}))g(\mu^t) = \lambda Q(\mu^t)P(\mu^t). \tag{9}$$

This result emphasizes that when heterogeneity across households with respect to the asset market participation is taken into account, the link between asset prices and aggregate consumption is broken. Instead, it is the marginal utility of households who participate in asset markets that is linked to asset prices. This result can be generalized to asset prices for different maturities. For example, a n-period bond that costs q_t^n in period t and pays \$1 in all states in period $t+n$ can be derived using the period zero nominal asset price as

$$q_t^n = \beta^n E_t \frac{U'(c_{At+n})}{U'(c_{At})} \frac{P_t}{P_{t+n}}. \tag{10}$$

where reference to the state μ^t has been suppressed. This asset price differs from the one that arises in a standard cash-in-advance in that it depends on the marginal utility of the asset holders $U'(c_A)$ rather than that of the representative agent.

The implications of this result for the relationship between relative prices and relative consumption across countries are revealed by a two-country economy version of the model extending on the work of Lucas (1982). In addition to the money growth shock μ and the idiosyncratic endowment shock y, households are subject to a shock to the distribution of their idiosyncratic endowment denoted by η. The history over the aggregate shock is given by $s^t = (\mu^t, \eta^t)$ and is distributed according to the probability density function $g(s^t)$. This shock serves as a mechanism to generate a type of money demand shock to be offset by an endogenous monetary policy. The result is a money growth process less persistent than nominal interest rates as observed in the data. For simplicity, households are assumed to demand only domestic goods which they purchase using domestic currency in the goods market. As in the one-country economy a fixed cost of γ must be paid for transferring cash between asset and goods markets and households in each country choose to transfer only domestic currency between the two markets. In the asset market however they can save by investing in bonds denominated in either currency. Relationship between bonds denominated in either currency is established through lack of arbitrage and hence it satisfies $q_i(s^t, s_{t+1}) = q_j(s^t, s_{t+1})e(s^t)/e(s^{t+1})$ where $q_i(s^t, s_{t+1})$ is the price for a one-period bond denominated in country i's currency and $q_j(s^t, s_{t+1})$ is the price for a one-period bond denominated in country j's currency. $e(s^t)$ is the nominal exchange rate in terms of country i's currency per country j's currency. This relationship also holds for period zero nominal asset prices so that $e(s^t) = e_0 Q_j(s^t)/Q_i(s^t)$. Defining real exchange rate as $r(s^t) = e(s^t)P_j(s^t)/P_i(s^t)$ and using equation (9) yields the relationship that links real exchange rates to a ratio of marginal utilities of households who participate in asset markets:

$$r(s^t) = e_0 \frac{U'(c_{Ajt})}{U'(c_{Ait})}. \tag{11}$$

This relationship constitutes the basis for our empirical model which we test using micro data on household consumption and financial asset holding. The underlying asset market structure provides a guide in choosing the relevant set of financial assets to define asset market participation.

3 Data

We use data on household consumption and financial assets from the Survey of Household Income and Wealth (SHIW) collected by the Bank of Italy to test the equilibrium condition of interest. The choice of dataset is motivated by several empirical issues. First, in order to test the relationship given in equation (2) we need comparable datasets for countries in consideration. Datasets that contain information on both consumption and financial assets at the household level are rare. And in case they existed, the issue of comparability in terms of the information provided and the sample design would be problematic. To avoid this problem we test this condition for twenty administrative regions of Italy. The level of detail provided by the SHIW allows such a strategy. Since the nominal exchange rate e_t between any two regions of Italy is equal to one, equation (2) reduces to

$$\frac{P_{jt}}{P_{it}} = \frac{U'(c_{Ajt})}{U'(c_{Ait})}. \tag{12}$$

As we detail in the next subsection, the household dataset provides information on other household characteristics which we use to shed light on the asset market participation decision.

3.1 Household Data

The Bank of Italy Survey of Household Income and Wealth (SHIW) is part of the Historical Database (HD) that contains information on Italian household budgets for the period 1977-2002. In particular, the dataset has information on:

- Individual characteristics and occupational status
- Consumption expenditure (durable and non-durable)
- Household financial asset and liabilities

- Different sources of income of household members (payroll and self-employment income, pensions, transfers, and property income)
- Properties lived in or owned by the household.

Detailed information and definitions for consumption expenditure and financial assets are provided in the Data Appendix A.

The surveys have undergone numerous changes over time. Until 1987, they consisted of annual cross-sections, with the exception of 1985. Thereafter, the surveys were conducted biannually, with 1998 replacing 1997. A panel component has also been added to facilitate longitudinal analysis. In total, the dataset contains 17 cross-sections, 7 of them including panel households. Our empirical work focuses mainly on the sample period 1982-2002 covering 11 cross-sections as the surveys started collecting information on household financial asset holdings in 1982. In what follows, we describe some basic characteristics of these financial asset holdings for Italian households.

3.1.1 Asset Market Participation Among Italian Households

The dataset provides information on the holdings of a rich array of financial assets which are detailed in the data appendix. Most of these financial assets, however, have observations for only a subset of the cross-sections spanning the sample period 1982-2002. In particular, the following categories have observations for the whole sample period 1982-2002:

- bank deposits in current accounts
- bank deposits in savings accounts
- BOTs (T-bills)
- other Italian government securities (CTEs, CTOs et al.)

- corporate bonds
- shares.

Table 1 summarizes the incidence of asset holdings over time for each of these categories as well as the total. The riskless asset reported in the last column refers to the sum of bank deposits in current accounts, bank deposits in savings account, BOTs (Treasury bills up to one year maturity), and other government securities. One of the striking features of these figures is the increase in total asset holding by households between 1984 and 1987 from 35% to 80%. This increase is evident in all types of financial assets with varying degrees. Table 2 presents incidence of asset holdings by age group. Household asset holdings display a hump shape over the life cycle with the exception of holdings of savings accounts which is stable. Table 3 documents the percentage of asset holders by educational qualification. For every type of financial asset listed, the incidence of asset holding is higher among more educated households. Table 4 reveals the significant differences between regions of Italy in terms of asset ownership. For all types of financial assets, asset ownership is much higher among households who reside in the North and much lower for households who reside in the South or the Islands. Households who reside in Central Italy are also in the center in terms of their asset holdings. The only exception is the saving accounts which are most common in Central Italy and more evenly held throughout Italy.

3.2 Price Data and Real Exchange Rates

To construct bilateral real exchange rates between regions we use the Consumer Price Index for Blue and White Collar Worker Households (FOI) collected by the Italian National Institute of Statistics (ISTAT). The main reason for working with this index is that FOI provides information on consumer prices at the regional and provincial level for a total of

21 regions and 113 provinces of Italy. To make it compatible with the household data on financial assets and consumption, we use 20 regions listed at the end of Data Appendix A for which the SHIW provides the ISTAT regional codes.

The real exchange rate between any two regions i and j at time t is equal to P_{jt}/P_{it}, where P_{jt} and P_{it} denote the price levels at time t in region j and region i respectively.

The dataset spans the period 1947-2003 and is in monthly frequency. An average price index for each year is also included in the dataset and is equal to the arithmetic average of the price indices for the twelve months in that year. Therefore, P_{it} refers to the average price index for year t for region i.

We also adjusted the price indices to take into account the changes in the base year according to which they are calculated. During the sample period we study, five different base years (1980, 1985, 1989, 1992, and 1995) are used by ISTAT to calculate price indices. Using the coefficients provided in the FOI dataset, we converted price indices so that they all reflect changes from the one base year 1980.

4 Econometric Specification

We use generalized method of moments (GMM) to test the equilibrium condition given by equation (2). The econometric method proposed by Hansen and Singleton (1982) to estimate and test nonlinear rational expectations models has been widely used by the empirical literature on consumption-based asset pricing models. The estimation strategy involves using the theoretical economic model to generate a family of orthogonality conditions. Then the parameters of the model are estimated by minimizing a criterion function constructed from these conditions.

In the context of GMM estimators for nonlinear models, the starting point is to specify an elementary zero function which depends on the vector of observed variables X_t, one of which must be endogenous, and on the k dimensional vector of parameters θ to be estimated. Our choice for this elementary zero function is

$$h_t(\theta, X_t) = \left[\Delta \frac{P_{jt}}{P_{it}} - \frac{\Delta U'(c_{Ajt})}{\Delta U'(c_{Ait})}\right] \tag{13}$$

This condition differs from the one in equation (12) in that it is in growth rates rather than in levels. If the relationship given by equation (12) holds in levels, then it also holds in growth rates. Furthermore, using the growth rate of consumption allows us to focus on households' intertemporal consumption choices conditional on their asset market participation status in each period. The role of this elementary zero function is the same as a residual in the estimation of a regression model. Therefore, the expectation of this elementary zero function must vanish if it is evaluated at the true value of the parameters, but not in general otherwise. Therefore, for a given elementary function $h_t(\theta, X_t)$ for observation t, this key property can be written as

$$\mathrm{E}\left[\Delta \frac{P_{jt}}{P_{it}} - \frac{\Delta U'(c_{Ajt})}{\Delta U'(c_{Ait})}\right] = 0 \tag{14}$$

where E is the expectation operator under the data generating process that belongs to the model. The theoretical moment conditions are constructed using an l dimensional vector of instruments z_{it-1} for region i as follows

$$\mathrm{E}\left[z_{it-1}\left(\Delta \frac{P_{jt}}{P_{it}} - \frac{\Delta U'(c_{Ajt})}{\Delta U'(c_{Ait})}\right)\right] = 0. \tag{15}$$

To complete the specification of the moment condition, we need to choose the form of the utility function. This also determines what parameters are to be estimated. Suppose that the utility function is time-separable with the period utility having the power form

$$U(c) = \begin{cases} c^{1-\alpha}/(1-\alpha) & \alpha > 0, \quad \alpha \neq 1; \\ \ln c & \alpha = 1 \end{cases}$$

where α is the risk aversion parameter. Using this functional form for the period utility in equation (3) and assuming equal, constant discount factors yields the following:

$$\mathrm{E}\Big[z_{it-1}\Big(\Delta\frac{P_{jt}}{P_{it}} - \Big(\frac{\Delta c_{Ait}}{\Delta c_{Ajt}}\Big)^{\alpha}\Big)\Big] = 0. \tag{16}$$

Equation (16) reveals that the parameter to be estimated is the risk aversion parameter α. By identifying a reference region j, and estimating equation (16) for pairs of regions, α can be estimated for each region i separately in a similar fashion to Head, Mattina, and Smith (2003). The set of instruments z is then chosen for region i as indicated by its subscript. Among the instruments commonly used in the literature are lagged endogenous variables such as the lagged consumption growth. Other possible instruments include a constant and lagged asset returns. Using $l > k$ instrumental variables to estimate equation (14) yields $l - k$ over-identifying restrictions which can be tested using the relevant test statistics.

We estimate the risk aversion parameter α by minimizing an appropriately chosen quadratic form based on the moment condition. The method of moments estimator of the function given on the left-hand-side of equation (15) is given by

$$f(\alpha) = \frac{1}{T}\sum_{t}^{T}\Big[z_{it-1}\Big(\Delta\frac{P_{jt}}{P_{it}} - \Big(\frac{\Delta c_{Ait}}{\Delta c_{Ajt}}\Big)^{\alpha}\Big)\Big] \tag{17}$$

and is used to construct the criterion function $Q(\alpha) = f(\alpha)'Wf(\alpha)$ where W is a symmetric, positive definite weighting matrix. Estimation employs iterated GMM, a procedure that starts from an initial estimate $\hat{\alpha}$ to obtain an estimate of the asymptotic covariance matrix of the estimator, which then is used to obtain a second-round estimate of α. This second-round estimate $\hat{\alpha}$ can then be used to obtain yet another estimate of the covariance matrix and so on. Even though, for a correctly specified model, all the estimators have the same asymptotic distribution regardless of how many iterations are carried out, performing more iterations often improves the finite-sample properties of the estimator. Therefore, as suggested by Hansen, Heaton and Yaron (1996), a continuously updated estimator that has been iterated to convergence may yield the best results.

Besides its computational advantages, the GMM criterion function has a main advantage for testing purposes. In particular, the minimized value of the GMM criterion function $Q(\alpha)$ provides the Hansen-Sargan over-identification test statistic which is asymptotically distributed as $\chi^2(l-k)$. When the number of moment restrictions is greater than the number of parameters so that $l > k$, there are $l-k$ remaining linearly independent moment conditions that are not set to zero but should be close to zero if the model restrictions are true. A Hansen-Sargan test may reject the null for various reasons. In the case of feasible GMM estimation, especially involving heteroscedasticity and autocorrelation consistent (HAC) covariance matrices, this may be due to the finite-sample distribution of the test statistic being substantially different from its asymptotic distribution. One way to improve the finite-sample properties of the estimator is to use an iterative procedure as explained above.

An important part of the econometric specification is to define and construct endogenous variables used in moment condition given by equation (16). In particular, defining consumption growth Δc_{Ait} of a group of asset holders in a consistent manner with the theoretical model's assumptions and by utilizing information on households characteristics from the data is central to our empirical work. The next section provides details on alternative definitions of asset market participants as well as the econometric method used to identify them from the data.

5 Consumption Growth

The moment condition given by equation (16) relates bilateral real exchange rates to relative consumption growth of asset market participants in regions i and j. Ideally, this consumption growth should be computed for households who participate in asset markets in adjacent time periods in consideration. Obviously, panel data containing observations on asset market participation status together with consumption over a period of time are required for such an empirical strategy. The dataset used in this study contains panel households only for a subset of all the available cross-sections. Furthermore, the panel data suffer from severe attrition over time. For example, if we consider all cross-sections with a panel component covering the period 1987-2002, there are only 44 households who participated in all 8 surveys. We address this issue by using two different definitions of consumption growth. First, we employ an approach proposed by Attanasio et al. (2002) to identify likely asset holders in adjacent time periods based on their *predicted* probability of holding assets. The authors use this approach to identify likely and non-likely shareholders from the U.K. Family Expenditure Survey for the period 1978-1995 covering 18 consecutive years.

According to the definition used in this approach, the consumption growth for the group of likely asset holders in region i is given by the following:

$$\Delta(c_{Ait+1}) = [\overline{c_{it+1}} \mid p(y_{h,t+1} = 1)_t > p_t] - [\overline{c_{it}} \mid p(y_{h,t} = 1)_t > p_t]. \tag{18}$$

This grouping estimator identifies the household h as a likely asset holder if the predicted probability of owning assets at time t denoted by $p(y_h = 1)_t$ is greater than a cutoff point p_t. Then, $\overline{c_{it}}$ denotes the time t average consumption for the group of households who are identified to be likely asset holders at time t. The time $t+1$ average consumption for the group of households who are identified to be likely asset holders at time t is given by

$\overline{c_{it+1}}$. The subscript i refers to region $i = 1, \ldots, 20$. Predicted probabilities are computed using an estimated probit model of asset ownership. Our choice for the cutoff point p_t is the proportion of actual asset holders observed at time t. An important point to notice is that both terms on the right-hand side of the above equation are conditioned on the predicted probability of asset ownership at time t. Using the same criterion allows us to define the group of households over which average consumption growth is computed in a consistent manner. It should also be noted that $p(y_{h,t+1} = 1)_t$ in the first term of the right-hand-side of equation (18) is the predicted probability of asset ownership for households that are observed at time $t+1$ whereas $p(y_{h,t} = 1)_t$ in the second term is for households that are observed at time t. This specification obviously limits the choice of explanatory variables used in the probit estimation of asset ownership to those that are either perfectly predictable or constant over time.

Using this approach has the additional benefit of providing insight into factors that determine a household's asset market participation decision which itself is an interesting and active topic of research. Our second approach is to group households in each period on the basis of their current asset ownership. Then we compute consumption growth for this group of asset owners in adjacent periods. This specification provides a benchmark against which predictions and results from the first approach can be compared.

5.1 Probit Regression

In order to determine asset market participation status based on predicted probabilities, we run a probit regression using data on Italian households. The estimation sample is comprised of pooled data from 11 cross-sections for a total of 55,330 observations. Households with missing values for any of the financial assets (47 observations), non-durable consumption expenditure (75 observations), and educational qualification (11 observations) are again excluded. There are no restrictions on the basis of age or any other household characteristics placed on the estimation samples. Data on panel households, available from 1989 on, have not been exploited in either estimation sample in order to defer an explicit treatment of fixed effects to a later stage of the research.

5.1.1 Model

The probit model can be derived from the following latent variable model

$$y_h^* = X_h\beta + \epsilon_h, \qquad \epsilon_h \sim \mathrm{NID}(0,1)$$

with

$$y_h = \begin{cases} 1 & \text{if } y_h^* \geq 0 \\ 0 & \text{otherwise} \end{cases}$$

where y_h is a binary dependent variable that takes on a value of 1 if household h has holdings of the designated assets at the end of the year and 0 otherwise. Then, the probability of household h holding assets at the end of the period is given by $p(y_h = 1) = \Phi(X_h\beta)$ where $\Phi(\cdot)$ denotes the cumulative standard normal distribution function. X_h is the vector of explanatory variables and β is a vector of regression parameters.

In determining the set of financial assets used to define the dependent variable y_h, assumptions regarding the underlying asset market structure should be given careful consideration. The moment conditions given by equation (16) involve the consumption of households who trade a complete set of contingent claims subject to a fixed cost. We know that holding a complete set of contingent claims is equivalent to holding the riskless asset. Therefore, the price of a portfolio consisting of a complete set of contingent claims should be the same as the price of the riskless asset. Hence, defining the dependent variable for our probit regression by using the information on the holdings of the riskless assets available in the data should be a reasonable approximation. Among the financial assets surveyed during the 1982-2002 period, we choose bank deposits in current accounts, bank deposits in savings account, BOTs (Treasury bills up to one-year maturity), and other government securities to define the riskless asset. This definition is consistent with the one utilized by Guiso, Jappelli, and Terlizzese (1996) who study the role of income risk and borrowing constraints in portfolio choice using the 1989 cross-section of households from the SHIW. Their narrow definition of risky assets includes long-term government bonds (BTPs), corporate bonds, shares of mutual funds (investment fund units), and shares (equities). Safe assets are defined residually. For our sample period, this residual corresponds to the set of riskless assets we have listed above. The last columns of Tables 1-4 refer to the sum of the holdings of these riskless assets.

It is evident that the incidence of riskless asset holdings is very similar to the incidence of total asset holdings by time, age group, education, and geographical area of residence. The reason is that households who hold risky assets such as shares and corporate bonds tend to also hold less risky assets in their portfolios. For example, out of 2250 households who hold shares, 2214 also hold some type of riskless asset. Similarly, 1173 out of 1188 households who hold corporate bonds in their financial portfolio hold a riskless asset as well. In order to see how the results depend on the financial assets used to define asset market participation, we experiment with a larger set of assets as well. Accordingly we define the dependent variable to take on a value of 1 if household h has holdings of any of the assets listed in Table 1 at the end of the year and 0 otherwise.

For the explanatory variables, we use age, education, and time (trend) from the list of household characteristics available in the dataset. These variables are among the commonly used variables used in the related literature to explain asset holding behavior of households. Age is used as a proxy for life-cycle stage. We define it to be the current age of the household head which, in our estimation sample, takes on values between 16 and 114 and has a mean value of 52.26. Education could matter for asset market participation because informational costs associated with investing in assets are likely to be lower for educated households. We construct 4 educational qualification dummies: e1 for household heads with no or elementary education, e3 for household heads with high school education, and e4 for household heads with a bachelor's or a post-graduate degree. The omitted dummy e2 identifies households with middle school education. Trend is included to capture the changes in the incidence of asset ownership over time. An alternative specification includes regional dummies to take into account the significant regional variance in penetration of financial instruments in Italy. Three regional dummies identify households residing in North Italy, South Italy and the Islands, and Central Italy, with the last one being the omitted dummy.

Our baseline specification is similar to the one used by Attanasio et al. (2002) who study share ownership among U.K. households for the period 1978-1995. Their data source, the U.K. Family Expenditure Survey (FES) contains detailed information on the income and demographic characteristics of British households as well as expenditure patterns. Information on share holdings is imputed from responses to the income part of the questionnaire. They restrict their sample to households in which the head of the household is between ages 20 and 59. To construct education dummies, they divide households into three groups according to whether the household head left full-time education at or before the compulsory school-leaving age, between this age and 18, and above age 18. The penultimate group is referred to as having A levels while the last group is identified as having college education.

A common property of the explanatory variables used in probit regressions is that they are either perfectly predictable or constant over time with the exception of regional dummies. However, a casual observation of the panel component of the SHIW surveys suggests that geographical area of residence could be treated as constant from one period to the next. This property will be necessary to calculate the first term on the right-hand-side of equation (18).

5.2 Probit Estimation Results

Results presented in Tables 5-24 refer to two groups of regressions using a different set of financial assets to define the dependent variable. Computations of consumption growth for predicted asset holders utilize estimates from the probit regression presented in Table 8. This regression uses the set of riskless assets as explained earlier to define asset market participation. Results based on a broader set of assets are also provided for comparison. All computations are done using Stata version 7.0.

We use two different specifications starting with a probit regression that includes as explanatory variables higher orders of age, and time in addition to terms allowing for interactions between age and time, education and time as well as age and education. Results presented in Table 5 reveal that all coefficients are significant at the 95% confidence level except for the three terms allowing for interactions of the age variable with the no/elementary education dummy, with the bachelor's/post-graduate education dummy, and with the trend variable. The sign of the age coefficient is positive, while that of the squared age is negative. So there is a concave quadratic pattern of asset holding with age, disregarding all the age interaction terms. Dummies for high school education and bachelor's degree or post-graduate qualification have positive coefficients, while the dummy for no education or elementary education has a negative coefficient. Thus, apart from interaction terms, probability of asset ownership rises with level of education. This finding is not surprising given that education is an important factor in processing information regarding asset markets. These results are consistent with the findings of Attanasio et al. (2002). One difference worth noting between the two sets of results is the sign on the trend coefficients. Our trend variable has a negative coefficient which is followed by coefficients alternating in sign for higher orders of trend up to the fifth order. In Attanasio et al. (2002) these findings are reversed with a positive coefficient on trend and alternating through higher orders. It is not readily obvious what drives these results in our data so we experiment with alternative specifications that include only up to second, third, and fourth orders of the trend variable. The specification with up to fourth order of trend variable yields positive coefficients for all except for the third order. Other estimations produce a positive coefficient on the trend, a negative and insignificant coefficient on the squared trend, and a positive coefficient on the third order.

Next, we estimate a probit model with an alternative specification including the trend variable only up to second order. This parsimonious specification lends itself to a more straightforward interpretation of trend variables. We also use regional dummies with the omitted dummy a2 identifying households residing in Central Italy. Residents living in North Italy are identified with dummy a1 and residents living in the South and the Islands are identified with dummy a3. Results reported in Table 8 are similar for the variables included in both specifications such as age and education dummies. The main difference is that now trend has a positive coefficient while trend squared has a negative one. So interactions aside, there has been an increasing trend in asset ownership over time. The regional dummies are both significant with a positive coefficient for the North and a negative coefficient for the South and the Islands. Being from the North increases the probability of holding assets by 0.086 and being from the South or the Islands decreases it by 0.224 which is the largest marginal effect reported in Table 8.

Our results reveal that using a broader set of financial assets to define asset market participation yield similar findings under both the baseline specification and the alternative specification as shown in Tables 12 and 18 respectively. This is likely due to the fact that, in our sample, households that hold any one of the assets included in the broader set usually hold a riskless asset as well. We also run logit regressions on the broader set of assets as a robustness check. The t-ratios and marginal effects from the logit regressions are given in Tables 15 and 21 and are similar to those from the probit regression. According to a proportionality result by Amemiya (1981), coefficients of the logit model can be approximated by scaling coefficients of the probit regression by 1.6. Our logit regression yields parameter estimates that are about 1.7 times as large the probit estimates which is expected for unbalanced samples such as ours as noted by Greene (1993).

5.2.1 Marginal Effects

In probit analysis, marginal effects provide more insight than the simple parameter estimates since the latter are not readily interpretable in terms of the relevant variables' contribution to the probability of a positive outcome. To see this, remember that the probability of household h holding assets is given by $\Pr(y_h = 1) = \Phi(X_h\beta)$ where $\Phi(\cdot)$ denotes the cumulative standard normal distribution function. Therefore, as in Greene (1993), the marginal effects are given by

$$\begin{aligned} \frac{\partial E(y_h|X_h)}{\partial X} &= \left\{\frac{d\Phi(X_h\beta)}{d(X_h\beta)}\right\}\beta \\ &= \phi(X_h\beta)\beta \end{aligned}$$

where $\phi(\cdot)$ is the standard normal density. Obviously, the value of these marginal effects depends on the independent variables. One way of computing marginal effects is to evaluate these expressions at the sample means of the data. Alternatively, one can evaluate the marginal effects at every observation and use the sample average of the individual marginal effects. For large samples, these two will give a similar answer. We use the first approach which is the default option in Stata. According to the results presented in Table 8, a marginal increase in age of the household head increases the household's probability of holding assets by 0.020. The effect of squared age is negative but quite negligible. As for the dummy variables, the marginal effect gives the change in the probability of a positive outcome for a discrete change of dummy variable from 0 to 1. For example, the probability of holding assets is higher by 0.127 for a household with a head who has a bachelor's degree or a post-graduate qualification whereas it decreases by 0.065 below the default education level for a household that has a head with no or elementary education.

The marginal effects for the interaction terms reveal that the biggest contributions among the interactions come from the interaction of the education dummies with trend. The interpretation of these interaction terms should be done with caution. Ai and Norton (2003) point out that, when studying how the effect of one independent variable on the dependent variable depends on the magnitude on another independent variable, attention is usually and incorrectly placed on the marginal effect of the interaction term. This marginal effect, for example for the interaction term AgexTrend is given by $\partial\Phi(\cdot)/\partial(\text{AgexTrend}) = \beta_{AgexTrend}\Phi'(\cdot)$. However, the interaction effect is the cross derivative of the expected value of y and is equal to

$$\frac{\partial^2\Phi(\cdot)}{\partial\text{Age}\partial\text{Trend}} = \beta_{AgexTrend}\Phi'(\cdot) + (\beta_{Age} + \beta_{AgexTrend}\text{Trend})(\beta_{Trend} + \beta_{AgexTrend}\text{Age})\Phi''(\cdot).$$

Obviously, the two are not equal unless the second term in the interaction effect is equal to 0. As Ai and Norton (2003) suggest, the literature's focus on marginal effects when reporting estimation results for interaction terms might be due to the fact that statistical software packages like Stata compute these effects rather than the interaction effects.

Similar considerations apply to the interpretations of marginal effects for the explanatory variables that enter into regressions in higher orders. For example, our baseline regression includes among explanatory variables both age and age squared and therefore the appropriate marginal effect for age should be calculated as:

$$\frac{\partial E(y_h|X_h)}{\partial\text{Age}} = \left\{\frac{d\Phi(X_h\beta)}{d(X_h\beta)}\right\}[\beta_{Age} + 2\beta_{Age^2}\overline{\text{Age}}]$$

where $\overline{\text{Age}}$ is the sample mean for age. To provide more insight into the marginal effects of such terms, we run a stripped-down version of the probit regression without any interaction or higher order terms. The explanatory variables are age, educational dummies, regional dummies, and trend. We also run a separate regression for each year in our estimation sample. Results presented in Table 11 reveal that age yields a negative marginal effect for every year as well as the pooled data, though this represents the net effect of the concave quadratic age holding pattern. Age effects for years 1983, 1995, and 2000 turn out to be insignificant. Marginal effects for all other variables have the same sign as in the baseline and the alternative probit regression. Household heads with high school education or with a bachelor's degree and/or post-graduate qualification are more likely to hold assets while household heads with no education or an elementary education have a lower probability of holding assets. All variables are significant with the exception of dummy for high school education in year 1984.

5.3 Predicted Probabilities of Asset Ownership

We use the estimated coefficients from the probit regression presented in Table 8 to calculate the predicted probabilities of asset ownership. These probabilities are then used to construct the group of likely asset holders in accordance with equation (18). We start with the second term on the right-hand-side of the equation. $p(\text{assetholder})_t$ in this term refers to the predicted probability of asset ownership at time t for households who are observed at time t. We create dummies d1982, ...,d2000 that take on a value of 1 if this predicted probability is greater than the actual proportion of asset holders for time t cross-section and 0 otherwise. Table 9 documents the incidence of predicted asset ownership according to this criteria together with actual ownership. Next, we calculate the average consumption of households for whom this dummy is equal to 1 for each region. This gives us the $\overline{c_{Ait}}$ for $i = 1, \ldots, 19$. Next, using the estimated coefficients from the probit estimation and the time t values of explanatory variables, we calculate the predicted probabilities of asset ownership at time t for households observed at time $t+1$. This gives us the $p(\text{assetholder})_t$ in the first term on the right-hand-side of equation (18). Note that using variables that are perfectly predictable or constant over time allows us to calculate these probabilities since otherwise the lack of panel data would make it impossible to infer time t values of the variables for households that are observed at time $t+1$. Then, we create dummies e1983, ...,e2002 that take on a value of 1 if this predicted probability is greater than the actual proportion of asset holders at time t and 0 otherwise. We obtain $\overline{c_{Ait+1}}$ for $i = 1, \ldots, 19$ by calculating the average consumption of households for whom the dummy is equal to 1 for each region. Then, we calculate the consumption growth for the group of likely asset holders, $\Delta(c_{Ait+1})$ for $i = 1, \ldots, 19$ as given in equation (18).

Tables following each set of regression results show how well the grouping estimator did in predicting the likely and the unlikely asset holders. The first set of tables summarizes this information for dummies d1982, ...,d2000. Therefore, they show how well the particular specification of probit models did in predicting the actual asset holders observed at time t to be likely asset holders at time t. A comparison of the top left cell of Tables 6 and 9 reveal that the alternative specification does a better job than the baseline specification in this regard with 67% versus 59%. The second set of tables show the dummies e1983, ...,e2002, and hence the information presented in these tables reveals what percentage of the households observed at time $t+1$ are predicted to be likely and unlikely asset holders at time t. Again, the alternative specification has higher predictive power than the baseline specification as shown in the bottom left cell of Tables 7 and 10.

The regressions utilizing a larger set of assets (any one of the assets listed in Table 1) define asset market participation yield similar results. A higher percentage of asset holders at time t are predicted to be likely asset holders under the alternative specification than the baseline specification: 66% versus 59%. We also present results from the logit regressions. Overall, the probit and the logit models have similar predictive power for both the baseline specification and the alternative one. The logit model also predict a higher percentage of actual asset holders to be likely asset holders using the alternative specification; 67% as opposed to 60% using the baseline specification. As for the dummies e1983, ...,e2002, once again the alternative specification outperforms the baseline specification for both the probit and the logit regression, though with a smaller margin.

6 GMM estimation

We estimate moment conditions given by equation (16) for pairs of regions with Lazio acting as the reference region j. This allows for separate, region-specific preference parameters α. The Val d'Aosta region is excluded because there are very few observations on asset market participants for some cross-sections and none for others. We use two different sets of instrumental variables. The first set is comprised of a constant and lagged relative consumption growth. The second set includes in addition the regional nominal interest rates.

Tables 25 and 26 present results from the nonlinear GMM estimations utilizing the first and the second set of instruments respectively. The first two columns in the tables refer to the estimations of the moment conditions for predicted asset holders. The next two refer to the results for the actual asset holders.

One formal test of the underlying moment conditions is based on overidentification. We estimate parameter α first using the small set of instruments and hence with two moment restrictions. In this case, as Hansen (1982) shows, T, the number of observations times the minimized value of the criterion function $Q(\alpha)$ is asymptotically $\chi^2(2-1)$ under the null hypothesis that the moment restrictions of the model hold. These test statistics, also called the J-statistic, are provided in second and fourth column of Table 25 together with their p-value. According to our findings, the model restrictions are not rejected at the 5% level of significance for any of the regions except for the Molise region with Campobasso as its major city. When we estimate the model parameter with three instruments, the results in Table 26 reveal that the moments restrictions are accepted for all eighteen regions. These findings are robust to whether the moments conditions are estimated using the average consumption growth for the predicted or the actual asset holders.

To provide a comparison for these results, we also estimate moment restrictions replacing $\Delta(c_{Ait})$ with $\Delta(c_{it})$. The latter is the growth rate of aggregate consumption for region i averaged over households residing in that region. Hence, moment conditions for this estimation are in effect based on the relationship given by equation (1) relating bilateral real exchange rates to relative aggregate consumption across regions. Again, we use the same two sets of instruments as in our earlier estimations. Results are presented in Tables 27 and 28. We focus on the estimations that utilize the larger set of instruments and compare J-statistics reported in the second column of Tables 26 and 28. Our findings reveal that when the model restrictions are estimated using the aggregate consumption growth J-statistics turn out to larger with smaller p-values for 14 out of 18 regions than those yielded under the limited asset market participation. Although the evidence does not reveal rejections under the aggregate data, it nevertheless points to a potential for improving the model's implications by accounting for the heterogeneity in asset market participation status across households.

An analysis of our GMM estimates of the risk aversion parameter α and how they compare to other findings in the literature provide another measure of how well the model did in matching the empirical evidence. We report these estimates together with their standard errors in Tables 25-28. First of all, α should assume positive values to be consistent with the utility theory. As earlier, we focus on estimations that utilize the larger set of instruments as reported in Table 26. When estimations are run on predicted asset holders, for six out of eighteen regions the parameter estimates of α turn out to be positive and statistically significant at 5% level of significance. These parameter estimates are denoted with a star in Table 26. Three of these regions are from the South and the Islands, two of them are from North Italy and one of them is from Central Italy. The reference region Lazio is also located in Central Italy with its major city being Rome. Among five parameter estimates that are negative only one is statistically significant. When relative consumption growth is computed over actual asset holders, estimates of α are positive and statistically significant for three regions. This finding might be due to issues related to grouping households on the basis of their current asset ownership in each period as we did with actual asset holders. The composition of the group so defined is subject to changes over time. Since these changes imply that the group of asset holders in each period is defined over different households, changes in their average consumption are hard to interpret.

In order to see how our estimates of the risk aversion parameter fit in with other estimates reported in the literature, we start with findings presented by Hansen and Singleton (1982). They estimate parameter α using Euler equations of a nonlinear model of stock market returns. Their estimates range from 0.3502 to 0.9993 depending on the type of return used. However, it should be noted that their Euler equations involve measures of aggregate consumption. Since our approach emphasizes the use of micro data instead of aggregate data when the model economy is characterized by heterogeneity across households their estimates might not provide an appropriate comparison.

As we have mentioned earlier, there is a growing literature that looks at implications of limited asset market participation for households' intertemporal consumption choices. For example, Attanasio et al. (2002) estimate Euler equations based on a consumption-based capital asset pricing model using micro data on British households that control for nonparticipation. The estimates of the risk aversion parameter vary from 0.548 to 3.527 for actual asset holders depending on the rate of return used. Estimates obtained from the Euler equation for the Treasury bills are lower than those obtained from the Euler equation on share returns. Using predicted asset holders yield estimates ranging from 0.647 to 3.175. Again, estimations that utilize returns on Treasury bills yield lower estimates. When estimations are done on Euler equations for non-shareholders estimates vary from -27.058 to 20.524. However, these implausible values are not unexpected since Euler equations should not hold for households that do not hold the underlying assets.

Vissing-Jørgensen (2002) also studies the implications of limited asset market participation although she focuses on another parameter, the elasticity of intertemporal substitution. However, if preferences have the constant relative risk aversion (CRRA) form, the coefficient of relative risk aversion is equal to the inverse of the elasticity of intertemporal substitution. Since her estimates are obtained from a model with this type of preferences, we consider their implication for the value of the risk aversion parameter. The estimates of the elasticity of intertemporal substitution are around 0.3-0.4 for stock holders and around 0.8-1 for bond holders. The implied values of the risk aversion parameter are thus around 2.5-3.3 for stock holders and 1-1.25 for bond holders.

The findings reported in the literature do not suggest a clear consensus on what the value of the risk aversion parameter is. Moreover, the value of the estimates vary depending on the type of assets used to define asset market participation and on whether the estimation employs aggregate or micro data. Our estimates are on the lower side of those comparable in terms of these two criteria.

Next we plan to look at what happens to the value of the estimates if we were to estimate a common parameter for all regions instead of allowing for region-specific parameter values. An obvious way to do this is by placing cross-equation restrictions when estimating moments conditions for all 18 pairs of regions.

7 Conclusion

This paper uses micro household data to investigate whether and to what extent accounting for heterogeneity in household asset market participation can explain the delink between real exchange rates and relative consumption across economies. It is important to acknowledge that the nature of the existing gap between predictions of the standard model and empirical evidence is likely to have many dimensions and be due to model's abstraction from frictions in various markets. By focusing on limited asset market participation we not only seek to unravel its implications for the puzzle but also shed light on an interesting phenomenon in its own right.

Our empirical work reflects the nature of this approach by starting off with identification of asset holders from micro household data. Estimation of a probit model reveals that factors such as age, education, geographical area of residence as well as time (trend) contribute significantly to household asset market participation decision.

Next, we undertake an empirical investigation of the relationship that relates bilateral real exchange rates to relative consumption of the asset market participants across countries. The GMM estimation of moment conditions provides evidence in favor of the model's implications. In particular, overidentification restrictions of the model are not rejected at 5% level of significance. The evidence with respect to the estimates of the risk version parameter α is more mixed yielding positive and statistically significant parameters for one third of the regions. However, results appear to be similar when moment conditions are based on aggregate consumption. A more refined and detailed study of asset market participation might produce a more clear cut distinction between the two sets of results. A natural step in this direction seems to be using a balanced panel approach when defining asset market participation. While our observations will be restricted to fewer cross sections we will be able to follow the asset market participation behavior of the same asset holders over a period of time. Using panel households will also allow us to have information with regards to not just their asset market participation status but also how their asset holdings change over time.

Another advantage to using a balanced panel is that we can investigate the role of participation costs in generating limited asset market participation by testing for structural state dependence. Structural state dependence refers to a situation where individuals who have experienced an event in the past are more likely to experience that event in the future, as a consequence of the event affecting future preferences, prices or constraints. In the context of the participation cost, the effects work through the budget constraints. According to Vissing-Jørgensen (2002), per period participation cost will not lead to structural state dependence in a household's portfolio decision. Instead, the importance of the per period participation cost can be determined by estimating the benefits of stock market participation for each household. This approach is similar in flavor to the theoretical framework used in Alvarez et al. (2002). In the model, the measure of the net gain from switching from being a nonparticipant to a participant is simple and static with only current variables. This simplicity stems from the assumption that the cash-in-advance constraint binds, so that a household's decision to pay the fixed cost in period t does not affect its real balances and consumption in future periods. Since current period's market participation decision has no effect on future period's budget constraint, structural state dependence is not at work in this context. Performing tests for structural state dependence in the participation decision would provide a justification, or lack thereof, for the underlying assumptions with respect to the economic environment that yields the equilibrium conditions we are ultimately interested in testing.

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A Data

A.1 Household Data

The household dataset is the Survey of Household Income and Wealth (SHIW) collected by the Bank of Italy. The dataset is part of the Historical Database (HD) that contains information on Italian household budgets for the period 1977-2002. In particular, the dataset has information on:

- Individual characteristics and occupational status
- Different sources of income of household members (payroll and self-employment income, pensions, transfers, and property income)
- Consumption expenditure (durable and non-durable)
- Properties lived in or owned by the household
- Household financial asset and liabilities.

The surveys have undergone numerous changes over time. Until 1987, they consisted of annual cross-sections, with the exception of 1985. Thereafter, the surveys were conducted biannually, with 1998 replacing 1997. A panel component has also been added to facilitate longitudinal analysis. In total, the dataset contains 17 cross-sections, 7 of them including panel households. The following table breaks down household participation in surveys between 1987-2002 into a panel component and new observations. For example, of the 8,011 households who have been surveyed in 2002 only 44 have participated since 1987 while 4,406 have been surveyed for the first time.

Households Interviewed in the 1987-2002 Surveys

	Year of Survey							
Year of First Interview	1987	1989	1991	1993	1995	1998	2000	2002
1987	8,027	1,206	350	173	126	85	61	44
1989		7,068	1,837	877	701	459	343	263
1991			6,001	2,420	1,752	1,169	832	613
1993				4,619	1,066	583	399	270
1995					4,490	373	245	177
1998						4,478	1,993	1,224
2000							4,128	1,014
2002								4,406
Sample Size	8,027	8,274	8,188	8,089	8,135	7,147	8,001	8,011

Some variables of interest are available only for a subset of these cross-sections. In particular, information on household financial assets is available starting in 1982 for the following categories:

- bank deposits in current accounts
- bank deposits in savings accounts
- BOTs (T-bills)
- other Italian government securities (CTEs, CTOs et al.)
- bonds
- shares.

A much more comprehensive set of financial assets is surveyed starting in 1987 and especially in 1989. The following table lists various categories of financial assets together with the time period they are available for. 1986 is excluded from this table as no data on financial assets were collected for this cross-section.

Financial Asset Categories

Year	1982	1983	1984	1987	1989	1991	1993	1995	1998	2000	2002
Bank deposits in CAs	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Bank deposits in SAs	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Certificates of deposits					$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Repos								$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Postal deposits in CAs					$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Postal deposits in SAs					$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Postal deposits in CAs/SAs					$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Postal savings certificates					$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Postal deposits/savings				$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
BOTs (T-bills)	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
CCTs (T-certificates)				$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
BTPs (T-bonds)					$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
CTZs (zero coupons)								$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Other gov't securities	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Bonds	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Shares of mutual funds				$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Shares	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Shares (limited liability)						$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Shares (partnerships)						$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Managed savings						$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Foreign securities								$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$
Loans to cooperatives								$\diamondsuit$	$\diamondsuit$	$\diamondsuit$	$\diamondsuit$

NOTE: CA refers to current account, SA refers to savings account.

Consumption is defined as the sum of expenditure on food consumption, entertainment, education, clothes, medical expenses, housing repairs and additions, and imputed rents. Expenditure on durable goods such as vehicles, furniture and appliances, or art objects are not included in the definition of consumption. Data on consumption expenditure are available for a total of 12 cross-sections starting with the 1980 survey. Again, the 1986 survey does not contain data on non-durable consumption.

Sampling for the survey is conducted in two stages. First, municipalities, which are the primary sampling units, are stratified by region and demographic size. Within each stratum, municipalities are randomly selected with a probability proportional to their size. In the second stage, the individual households to be interviewed are selected randomly.

To form the panel, municipalities were chosen from those already sampled in the previous survey. Households residing in these panel municipalities who had participated in at least two surveys were all included in the sample. The rest of the panel households are selected randomly from those who participated in the previous survey only.

A.2 Price Data

The price dataset is the Consumer Price Index for Blue and White Collar Worker Households (FOI). The Italian National Institute of Statistics (ISTAT) started constructing the dataset in 1947 under the name Index Numbers of Cost of Living in Provincial and Regional Capitals. As suggested by the current name which has been in effect since 1968, the price indices are calculated in accordance with the standard budgets of the blue and white collar worker households in extra agricultural sector. The dataset provides monthly price indices for major cities of 21 regions and 103 provinces. An annual price index is also calculated for each region as the arithmetic average of monthly indices in that year.

A.3 Interest Rate Data

Interest rate data are from the Bank of Italy statistical database available online at

http://bip.bancaditalia.it/4972unix/homebipeng.htm.

Data on interest rates are available starting in 1989. Due to lack of a single data source that spans the whole period to date, we construct the time-series for interest rates using the following three tables:

TDB20620 Total sight and time saving deposits by branch location (region) and size of deposit

TDC20640 Total deposits by branch location (region) and size of deposit

TDC20645 Total deposits by branch location (region) and size of deposit

The following table provides a list of 20 administrative regions and their major cities.

Regional codes IREG and PVDIP are for the Bank of Italy Survey of Household Income and Wealth and the Bank of Italy Monetary Statistics Survey respectively.

REGION	MAJOR CITY	IREG CODE	AREA	PVDIP
Piemonte	Torino	1	North	10010
Val d'Aosta	Aosta	2	North	10012
Lombardy	Milano	3	North	10016
Trentino-Alto Adige	Trento	4	North	10018
Veneto	Venezia	5	North	10020
Friuli-Venezia Giulia	Trieste	6	North	10022
Liguria	Genova	7	North	10014
Emilia-Romagna	Bologna	8	North	10024
Tuscany	Firenze	9	Centre	10028
Umbria	Perugia	10	Centre	10030
Marche	Ancona	11	Centre	10026
Lazio	Roma	12	Centre	10032
Abruzzo	L'aquila	13	South & Islands	10036
Molise	Campobasso	14	South & Islands	10038
Campania	Napoli	15	South & Islands	10034
Puglia	Bari	16	South & Islands	10040
Basilicata	Potenza	17	South & Islands	10042
Calabria	Reggio Calabria	18	South & Islands	10044
Sicily	Palermo	19	South & Islands	10046
Sardinia	Cagliari	20	South & Islands	10048

B Tables

B.1 Asset Holdings

Table 1

Percentage of Asset Holders Over Time

Year	T-bills	Corporate Bonds	Shares	Other Gov't Securities	Current Account	Savings Account	Total Assets	Riskless Assets
1982	3.15	0.20	0.66	0.50	17.50	20.40	36.22	35.94
1983	4.31	0.19	0.49	0.97	20.53	17.07	35.89	35.82
1984	4.72	0.17	0.41	0.79	19.56	17.26	35.31	35.23
1987	17.40	1.74	5.24	1.79	63.46	35.37	79.65	79.65
1989	18.93	1.08	4.40	0.90	63.87	35.16	81.49	81.46
1991	22.10	1.17	2.57	0.85	66.07	27.96	80.45	80.45
1993	18.43	2.30	3.82	0.82	66.62	29.13	79.72	79.70
1995	20.83	2.07	3.32	0.36	64.76	24.44	77.32	77.14
1998	8.82	4.67	7.35	0.22	74.41	22.78	81.87	81.76
2000	9.11	5.96	8.89	0.24	68.80	12.21	73.57	73.26
2002	7.88	5.15	6.49	0.41	72.24	11.33	76.67	76.60

Table 2

Percentage of Asset Holders by Age

Age Group	T-bills	Corporate Bonds	Shares	Other Gov't Securities	Current Account	Savings Account	Total Assets	Riskless Assets
≤ 30	8.26	1.38	2.22	0.66	60.02	22.26	70.19	70.17
31-40	11.55	1.96	4.42	0.62	65.13	23.14	74.82	74.71
41-50	14.08	2.49	5.38	0.81	64.14	25.51	74.75	74.65
51-65	16.07	2.50	4.61	1.14	56.71	25.41	70.78	70.71
≥ 65	12.31	1.74	2.39	0.54	40.20	24.91	58.55	58.45

Table 3

Percentage of Asset Holders by Education

Education	T-bills	Corporate Bonds	Shares	Other Gov't Securities	Current Account	Savings Account	Total Assets	Riskless Assets
None	3.62	0.20	0.18	0.15	17.08	23.83	37.82	37.80
Elementary	10.39	0.78	1.16	0.34	42.52	27.08	60.79	60.74
Middle school	12.56	1.72	2.75	0.71	61.95	24.27	73.21	73.10
High school	17.90	3.88	7.87	1.33	76.36	22.96	83.88	83.78
Bachelor's	26.45	6.33	13.49	2.61	82.32	21.52	88.72	88.46
Post-graduate	33.50	7.28	18.45	2.54	93.69	30.10	99.03	99.03

Table 4

Percentage of Asset Holders by Region

Region	T-bills	Corporate Bonds	Shares	Other Gov't Securities	Current Account	Savings Account	Total Assets	Riskless Assets
North	19.79	3.61	6.62	1.27	70.75	25.42	80.87	80.80
Central	13.06	1.65	3.06	0.60	58.53	26.13	73.50	73.37
South&Islands	5.51	0.59	1.45	0.34	36.64	22.81	52.68	52.60

B.2 Probit and Logit Regressions

Table 5

Probit Estimation for Asset Ownership

Baseline Specification (Riskless Asset)

Variable	Parameter	Standard Error	t-ratio	Marg. Effect
Age	.063	.003	21.77	.021
Age^2	-.001	.000	-23.03	-.000
e1 (None/Elementary)	-.433	.066	-6.58	-.144
e3 (High school)	.398	.068	5.80	.121
e4 (Bachelor's/Post graduate)	.537	.116	4.63	.149
Agexe1	.002	.001	1.38	.000
Agexe3	-.004	.001	-2.75	-.001
Agexe4	-.000	.002	-0.11	-.000
Trend	-.407	.047	-8.62	-.134
Trend^2	.209	.012	16.88	.069
Trend^3	-.026	.001	-18.76	-.009
Trend^4	.001	.000	18.51	.000
Trend^5	-.000	.000	-17.68	-.000
AgexTrend	-.000	.000	-1.54	-.000
AgexTrend^2	.000	.000	4.42	.000
e1xTrend	-.013	.002	-5.05	-.004
e3xTrend	.019	.003	6.59	.006
e4xTrend	.024	.005	4.96	.008
Constant	-1.570	.102	-15.40	
Observations	55,330			
Pseudo R^2	0.1899			
Log-likelihood	-27981.777			

Table 6

Incidence of Asset Ownership: Actual and Predicted

Probit Model (Riskless Asset)

Actual	Predicted	
	yes(d=1)	no(d=0)
yes (asset=1)	22,648 (59%)	15,894 (41%)
no (asset=0)	5,766 (34%)	11,123 (66%)

Table 7

Incidence of Asset Ownership: Actual and Predicted

Probit Model (Riskless Asset)

Actual	Predicted	
	yes(e=1)	no(e=0)
yes (asset=1)	22,352 (60%)	14,675 (40%)
no (asset=0)	4,343 (30%)	10,004 (70%)

Table 8

Probit Estimation for Asset Ownership

Alternative Specification (Riskless Asset)

Variable	Parameter	Standard Error	t-ratio	Marg. Effect
Age	.064	.003	21.31	.020
Age^2	-.001	.000	-21.69	-.000
e1 (None/Elementary)	-.205	.068	-3.00	-.065
e3 (High school)	.389	.071	5.47	.114
e4 (Bachelor's/Post graduate)	.464	.119	3.90	.127
Agexe1	-.001	.001	-0.81	-.000
Agexe3	-.003	.001	-1.98	-.001
Agexe4	.002	.002	0.89	.001
Trend	.383	.015	26.24	.122
$Trend^2$	-.016	.001	-25.57	-.005
AgexTrend	-.001	.000	-5.39	-.000
$AgexTrend^2$	.000	.000	8.36	.000
e1xTrend	-.016	.003	-5.95	-.005
e3xTrend	.016	.003	5.29	.005
e4xTrend	.025	.005	4.96	.008
North Italy	.274	.017	16.02	.086
South Italy & Islands	-.670	.017	-39.60	-.224
Constant	-2.234	.095	-23.37	
Observations	55,330			
Pseudo R^2	0.2432			
Log-likelihood	-25760.221			

Table 9

Incidence of Asset Ownership: Actual and Predicted

Probit Model (Riskless Asset)

Actual	Predicted	
	yes(d=1)	no(d=0)
yes (asset=1)	25,604 (67%)	12,848 (33%)
no (asset=0)	6,050 (36%)	10,839 (64%)

Table 10

Incidence of Asset Ownership: Actual and Predicted

Probit Model (Riskless Asset)

Actual	Predicted	
	yes(e=1)	no(e=0)
yes (asset=1)	25,989 (70%)	11,038 (30%)
no (asset=0)	4,750 (33%)	9,597 (67%)

Table 11

Marginal Effects

Probit Model with No Interaction Terms (Riskless Asset)

Marginal effects	Age	e1	e3	e4	a1	a3	t
Pooled data	-.002	-.121	.112	.183	.078	-.209	.020
1982	-.001	-.094	.091	.187	.072	-.149	
1983	-.001	-.046	.098	.314	.119	-.066	
1984	-.002	-.073	.035	.114	.049	-.123	
1987	-.002	-.124	.096	.156	.074	-.181	
1989	-.002	-.099	.101	.123	.054	-.172	
1991	-.002	-.112	.097	.135	.035	-.237	
1993	-.002	-.113	.108	.154	.069	-.189	
1995	-.001	-.107	.088	.152	.111	-.221	
1998	-.001	-.112	.090	.120	.050	-.230	
2000	-.001	-.122	.122	.188	.084	-.306	
2002	-.001	-.106	.119	.165	.098	-.260	

Table 12

Probit Estimation for Asset Ownership

Baseline Specification (Total Assets)

Variable	Parameter	Standard Error	t-ratio	Marg. Effect
Age	.064	.003	21.95	.021
Age^2	-.001	.000	-23.11	-.000
e1 (None/Elementary)	-.431	.066	-6.53	-.143
e3 (High school)	.406	.069	5.91	.123
e4 (Bachelor's/Post graduate)	.558	.116	4.79	.153
Agexe1	.001	.001	1.30	.000
Agexe3	-.004	.001	-2.92	-.001
Agexe4	-.001	.002	-0.22	-.000
Trend	-.418	.047	-8.81	-.137
$Trend^2$	.212	.012	17.07	.070
$Trend^3$	-.027	.001	-18.94	-.009
$Trend^4$	.001	.000	18.68	.000
$Trend^5$	-.000	.000	-17.86	-.000
AgexTrend	-.000	.000	-1.62	-.000
$AgexTrend^2$	.000	.000	4.49	.000
e1xTrend	-.013	.002	-5.12	-.004
e3xTrend	.019	.003	6.74	.006
e4xTrend	.025	.005	4.99	.008
Constant	-1.574	.102	-15.44	
Observations	55,330			
Pseudo R^2	0.1893			
Log-likelihood	-27561.252			

Table 13

Incidence of Asset Ownership: Actual and Predicted

Probit Model (Total Assets)

Actual	Predicted	
	yes(d=1)	no(d=0)
yes (asset=1)	20,626 (59%)	14,489 (41%)
no (asset=0)	5,500 (35%)	10,309 (65%)

Table 14

Incidence of Asset Ownership: Actual and Predicted

Probit Model (Total Assets)

Actual	Predicted	
	yes(e=1)	no(e=0)
yes (asset=1)	19,621 (59%)	13,697 (41%)
no (asset=0)	4,581 (34%)	9,060 (66%)

Table 15

Logit Estimation for Asset Ownership

Baseline Specification (Total Assets)

Variable	Parameter	Standard Error	t-ratio	Marg. Effect
Age	.110	.005	21.94	.021
Age^2	-.001	.000	-23.03	-.000
e1 (None/Elementary)	-.667	.117	-6.58	-.143
e3 (High school)	.616	.127	5.22	.110
e4 (Bachelor's/Post graduate)	.786	.225	3.59	.123
Agexe1	.003	.002	1.54	.000
Agexe3	-.007	.002	-2.70	-.001
Agexe4	-.001	.004	0.11	.000
Trend	-.144	.092	-9.47	-.143
Trend^2	.616	.026	17.50	.070
Trend^3	-.081	.003	-19.15	-.009
Trend^4	.004	.000	18.75	.000
Trend^5	-.000	.000	-17.83	-.000
AgexTrend	-.001	.000	-1.78	-.000
AgexTrend^2	.000	.000	4.77	.000
e1xTrend	-.033	.005	-5.65	.005
e3xTrend	.049	.006	8.06	.008
e4xTrend	.082	.012	6.58	.013
Constant	-2.067	.182	-15.03	
Observations	55,330			
Pseudo R^2	0.1890			
Log-likelihood	-27571.748			

Table 16

Incidence of Asset Ownership: Actual and Predicted

Logit Model (Total Assets)

Actual	Predicted	
	yes(d=1)	no(d=0)
yes (asset=1)	21,011 (60%)	14,079 (40%)
no (asset=0)	4,993 (32%)	10,844 (68%)

Table 17

Incidence of Asset Ownership: Actual and Predicted

Logit Model (Total Assets)

Actual	Predicted	
	yes(e=1)	no(e=0)
yes (asset=1)	20,194 (60%)	13,465 (40%)
no (asset=0)	4,009 (30%)	9,291 (70%)

Table 18

Probit Estimation for Asset Ownership

Alternative Specification (Total Assets)

Variable	Parameter	Standard Error	t-ratio	Marg. Effect
Age	.065	.003	21.50	.020
Age^2	-.001	.000	-21.80	-.000
e1 (None/Elementary)	-.200	.068	-2.93	-.064
e3 (High school)	.397	.071	5.58	.116
e4 (Bachelor's/Post graduate)	.486	.120	4.06	.131
Agexe1	-.001	.001	-0.91	-.000
Agexe3	-.003	.001	-2.15	-.001
Agexe4	.001	.002	0.77	.000
Trend	.383	.015	26.23	.121
$Trend^2$	-.016	.001	-25.53	-.005
AgexTrend	-.001	.000	-5.43	-.000
$AgexTrend^2$	.000	.000	8.39	.000
e1xTrend	-.016	.003	-6.04	-.005
e3xTrend	.016	.003	5.46	.005
e4xTrend	.026	.005	5.03	.008
North Italy	.273	.017	15.93	.085
South Italy & Islands	-.224	.006	-39.74	-.224
Constant	-2.250	.096	-23.52	
Observations	55,330			
Pseudo R^2	0.2444			
Log-likelihood	-25689.701			

Table 19

Incidence of Asset Ownership: Actual and Predicted

Probit Model (Total Assets)

Actual	Predicted	
	yes(d=1)	no(d=0)
yes (asset=1)	23,233 (66%)	11,854 (34%)
no (asset=0)	5,474 (35%)	10,363 (65%)

Table 20

Incidence of Asset Ownership: Actual and Predicted

Probit Model (Total Assets)

Actual	Predicted	
	yes(e=1)	no(e=0)
yes (asset=1)	20,758 (62%)	12,901 (38%)
no (asset=0)	3,629 (27%)	9,671 (73%)

Table 21

Logit Estimation for Asset Ownership

Alternative Specification (Total Assets)

Variable	Parameter	Standard Error	t-ratio	Marg. Effect
Age	.111	.005	21.32	.020
Age^2	-.001	.000	-21.48	-.000
e1 (None/Elementary)	-.349	.117	-2.99	-.064
e3 (High school)	.664	.126	5.27	.109
e4 (Bachelor's/Post graduate)	.753	.222	3.39	.114
Agexe1	-.001	.002	-0.69	-.000
Agexe3	-.005	.004	-2.14	-.001
Agexe4	.004	.004	0.80	.001
Trend	.663	.026	25.88	.120
$Trend^2$	-.028	.001	-25.31	-.005
AgexTrend	-.003	.000	-5.75	-.000
$AgexTrend^2$	.000	.000	8.74	.000
e1xTrend	-.029	.004	-6.41	-.005
e3xTrend	.035	.005	6.56	.006
e4xTrend	.066	.011	6.16	.012
North Italy	.468	.030	15.74	.083
South Italy & Islands	-.386	.010	-39.53	-.070
Constant	-3.875	.165	-23.50	
Observations	55,330			
Pseudo R^2	0.2443			
Log-likelihood	-25691.708			

Table 22

Incidence of Asset Ownership: Actual and Predicted

Logit Model (Total Assets)

Actual	Predicted	
	yes(d=1)	no(d=0)
yes (asset=1)	23,618 (67%)	11,469 (33%)
no (asset=0)	5,380 (34%)	10,457 (66%)

Table 23

Incidence of Asset Ownership: Actual and Predicted

Logit Model (Total Assets)

Actual	Predicted	
	yes(e=1)	no(e=0)
yes (asset=1)	20,931 (62%)	12,728 (38%)
no (asset=0)	3,590 (27%)	9,710 (73%)

Table 24

Marginal Effects

Probit Model with No Interaction Terms (Total Assets)

Marginal effects	Age	e1	e3	e4	a1	a3	t
Pooled data	-.002	-.121	.112	.185	.078	-.209	.020
1982	-.001	-.097	.088	.202	.071	-.151	
1983	-.001	-.047	.098	.316	.118	-.068	
1984	-.002	-.072	.036	.117	.049	-.124	
1987	-.002	-.124	.096	.156	.074	-.181	
1989	-.002	-.010	.101	.123	.054	-.172	
1991	-.002	-.112	.097	.135	.035	-.237	
1993	-.002	-.114	.107	.154	.068	-.190	
1995	-.001	-.107	.090	.154	.109	-.222	
1998	-.001	-.112	.090	.118	.044	-.229	
2000	-.001	-.123	.124	.193	.081	-.308	
2002	-.001	-.108	.117	.164	.099	-.259	

B.3 Nonlinear GMM Estimation

Table 25: GMM Estimation: Sample of Asset Market Participants (2 instruments, df=1)

	Predicted		Actual	
Region	$\hat{\alpha}$ (s.e.)	χ^2 (p-value)	$\hat{\alpha}$ (s.e.)	χ^2 (p-value)
Piemonte	-0.050	0.122	-0.011	1.051
	(0.064)	(0.73)	(0.079)	(0.30)
Lombardy	-0.057	0.348	-0.015	0.476
	(0.099)	(0.35)	(0.089)	(0.49)
Trentino-Alto Adige	-0.023	0.713	0.008	0.548
	(0.035)	(0.40)	(0.035)	(0.46)
Veneto	-0.542	1.425	-0.087	1.619
	(0.101)	(0.75)	(0.097)	(0.20)
Friuli-Venezia Giulia	0.092	1.839	-0.019	0.217
	(0.279)	(0.17)	(0.058)	(0.64)
Liguria	0.001	0.006	0.021	0.027
	(0.048)	(0.93)	(0.037)	(0.87)
Emilia-Romagna	0.017	0.131	0.051	0.901
	(0.083)	(0.72)	(0.112)	(0.34)
Tuscany	0.016	0.103	-0.015	0.304
	(0.077)	(0.75)	(0.033)	(0.58)
Umbria	0.019	2.546	0.015	2.404
	(0.042)	(0.11)	(0.050)	(0.12)
Marche	-0.125	0.006	-0.075	0.605
	(0.265)	(0.94)	(0.141)	(0.44)
Abruzzo	0.072	0.672	0.093	0.545
	(0.059)	(0.41)	(0.096)	(0.46)
Molise	-0.014	3.903	0.052	1.374
	(0.024)	(0.05)	(0.046)	(0.24)
Campania	-0.019	0.002	-0.004	1.321
	(0.075)	(0.96)	(0.149)	(0.25)
Puglia	-0.216	0.225	-0.051	0.646
	(0.370)	(0.63)	(0.181)	(0.42)
Basilicata	-0.024	1.585	-0.102	1.721
	(0.014)	(0.21)	(0.106)	(0.19)
Calabria	0.084	0.009	0.055	0.014
	(0.194)	(0.92)	(0.048)	(0.91)
Sicily	0.038	0.779	0.161	0.007
	(0.045)	(0.38)	(0.090)	(0.93)
Sardinia	0.124	0.265	0.173	0.91
	(0.210)	(0.61)	(0.176)	(0.34)

Table 26: GMM Estimation: Sample of Asset Market Participants (3 instruments, df=2)

	Predicted		Actual	
Region	$\hat{\alpha}$ (s.e.)	χ^2 (p-value)	$\hat{\alpha}$ (s.e.)	χ^2 (p-value)
Piemonte	-0.058	0.350	-0.090	0.984
	(0.041)	(0.83)	(0.059)	(0.61)
Lombardy	0.038	1.196	0.122	0.533
	(0.119)	(0.55)	(0.264)	(0.76)
Trentino-Alto Adige	0.012	0.602	0.007	1.045
	(0.019)	(0.74)	(0.016)	(0.59)
Veneto	-0.101	0.836	-0.090	0.882
	(0.004)	(0.66)	(0.045)	(0.64
Friuli-Venezia Giulia	0.150*	0.292	0.095	0.361
	(0.075)	(0.86)	(0.070)	(0.83)
Liguria	0.050*	1.514	0.067	1.785
	(0.015)	(0.47)	(0.038)	(0.41)
Emilia-Romagna	-0.048	0.051	-0.052	1.543
	(0.027)	(0.97)	(0.062)	(0.46)
Tuscany	0.060	0.593	0.076	0.832
	(0.095)	(0.74)	(0.087)	(0.66)
Umbria	0.030*	1.941	0.035*	2.396
	(0.002)	(0.38)	(0.002)	(0.30)
Marche	0.026	1.251	0.025	2.464
	(0.033)	(0.53)	(0.034)	(0.29)
Abruzzo	0.172	0.118	0.094	1.029
	(0.149)	(0.94)	(0.078)	(0.59)
Molise	0.016*	1.492	0.014*	1.878
	(0.008)	(0.47)	(0.004)	(0.39)
Campania	0.202	0.315	0.517	0.023
	(0.202)	(0.85)	(0.330)	(0.98)
Puglia	0.258	0.187	0.218*	0.880
	(0.223)	(0.91)	(0.078)	(0.64)
Basilicata	-0.368	0.513	0.453	0.949
	(0.529)	(0.77)	(0.299)	(0.62)
Calabria	-0.002	1.860	0.098	0.188
	(0.036)	(0.39)	(0.057)	(0.91)
Sicily	0.128*	0.665	0.112*	0.322
	(0.031)	(0.72)	(0.034)	(0.85)
Sardinia	0.100*	0.881	0.218	0.528
	(0.033)	(0.64)	(0.191)	(0.77)

Table 27: GMM Estimation of Moment Conditions:
Whole Sample (2 instruments, df=1)

Region	$\hat{\alpha}$ (s.e.)	χ^2 (p-value)
Piemonte	-0.005	0.951
	(0.010)	(0.33)
Lombardy	-0.005	0.482
	(0.063)	(0.49)
Trentino-Alto Adige	-0.000	0.597
	(0.023)	(0.44)
Veneto	0.014	2.821
	(0.073)	(0.09)
Friuli-Venezia Giulia	0.065	0.793
	(0.062)	(0.37)
Liguria	0.008	0.010
	(0.040)	(0.92)
Emilia-Romagna	0.058	0.456
	(0.086)	(0.50)
Tuscany	-0.028	0.501
	(0.027)	(0.48)
Umbria	0.049	1.462
	(0.072)	(0.23)
Marche	-0.073	0.089
	(0.106)	(0.76)
Abruzzo	0.055	0.558
	(0.058)	(0.45)
Molise	0.008	3.794
	(0.026)	(0.05)
Campania	0.033	0.983
	(0.167)	(0.32)
Puglia	-0.064	0.146
	(0.477)	(0.70)
Basilicata	-0.028	2.228
	(0.050)	(0.13)
Calabria	0.084	0.011
	(0.117)	(0.91)
Sicily	0.202	0.018
	(0.148)	(0.89)
Sardinia	0.458	0.135
	(0.734)	(0.71)

Table 28: GMM Estimation of Moment Conditions:
Whole Sample (3 instruments, df=2)

Region	$\hat{\alpha}$ (s.e.)	χ^2 (p-value)
Piemonte	-0.053	0.015
	(0.037)	(0.99)
Lombardy	0.035	1.120
	(0.106)	(0.57)
Trentino-Alto Adige	-0.022	2.265
	(0.010)	(0.32)
Veneto	-0.101	0.865
	(0.070)	(0.65)
Friuli-Venezia Giulia	0.049	0.377
	(0.038)	(0.83)
Liguria	0.055*	1.578
	(0.018)	(0.45)
Emilia-Romagna	-0.138	0.259
	(0.054)	(0.88)
Tuscany	0.050	0.763
	(0.081)	(0.68)
Umbria	0.042*	1.383
	(0.012)	(0.50)
Marche	0.012	3.492
	(0.027)	(0.17)
Abruzzo	0.068	0.128
	(0.054)	(0.94)
Molise	0.012*	1.913
	(0.005)	(0.38)
Campania	0.338	0.647
	(0.304)	(0.72)
Puglia	0.244	3.562
	(0.591)	(0.17)
Basilicata	0.097	2.372
	(0.119)	(0.30)
Calabria	-0.000	2.241
	(0.065)	(0.33)
Sicily	0.081*	0.691
	(0.009)	(0.71)
Sardinia	0.196*	1.634
	(0.061)	(0.44)