CONSUMPTION INEQUALITY AND PARTIAL INSURANCE

Richard Blundell
Luigi Pistaferri
Ian Preston

THE INSTITUTE FOR FISCAL STUDIES
WP04/28

CONSUMPTION INEQUALITY AND PARTIAL INSURANCE*

Richard Blundell
University College London and Institute for Fiscal Studies.

Luigi Pistaferri
Stanford University and CEPR.

Ian Preston
University College London and Institute for Fiscal Studies.

May 2005

Abstract

This paper examines the transmission of income inequality into consumption inequality and in so doing investigates the degree of insurance to income shocks. Panel data on income from the PSID is combined with consumption data from repeated CEX cross-sections to identify the degree of insurance to permanent and transitory shocks. In the process we also present new evidence of the growth in the variance of permanent and transitory shocks in the US during the 1980s. We find some partial insurance of permanent income shocks with more insurance possibilities for the college educated and those nearing retirement. We find little evidence against full insurance for transitory income shocks except among low income households. Tax and welfare benefits are found to play an important role in insuring permanent shocks. Adding durable expenditures to the consumption measure suggests that durable replacement is an important insurance mechanism, especially for transitory income shocks.

Key words: Consumption, Insurance, Inequality.

JEL Classification: D52; D91; I30.

*We would like to thank Joe Altonji, Orazio Attanasio, David Johnson, Arie Kapteyn, John Kennan, Robert Lalonde, Hamish Low, Bruce Meyer, Samuel Pienknagura, and Ken West for helpful comments. Thanks are also due to Erich Battistin for providing the CEX data, and Cristobal Huneeus for able research assistance. The paper is part of the program of research of the ESRC Centre for the Microeconomic Analysis of Public Policy at IFS. Financial support from the ESRC, the Joint Center for Poverty Research/Department of Health and Human Services and the National Science Foundation (under grant SES-0214491) is gratefully acknowledged. All errors are ours.

1 Introduction

Under complete markets agents can sign contingent contracts providing full insurance against idiosyncratic shocks to income. Moral hazard and asymmetric information, however, make these contracts hard to implement, and in fact they are rarely observed in reality. Even a cursory look at consumption and income data reveals the weakness of the complete markets hypothesis. Thus volatility of individual consumption is much higher than the volatility of aggregate consumption, a fact against full insurance [Aiyagari, 1994]. Moreover, there is a substantial amount of mobility in consumption [Jappelli and Pistaferri, 2004]. Formal tests of the complete markets hypothesis [see Attanasio and Davis, 1996], have often found the null hypothesis of full consumption insurance is rejected. Attempts to salvage the theory by allowing for risk sharing within the family and no risk sharing among unrelated families have also been unable to find evidence of complete insurance [Hayashi, Altonji and Kotlikoff, 1996].

In the textbook permanent income hypothesis the only mechanism available to agents to smooth income shocks is personal savings. The main idea is that people attempt to keep the expected marginal utility of consumption stable over time. Since insurance markets for income fluctuations are assumed to be absent, the marginal utility of consumption is not stabilized across states. If income is shifted by permanent and transitory shocks, self-insurance through borrowing and saving may allow intertemporal consumption smoothing against the latter but not against the former [Deaton, 1992]. This is simply because one cannot borrow to smooth out a permanent income decline without violating the budget constraint, so that permanent shocks to income will be permanent shocks to consumption.[1]

Models that feature complete markets and those that allow for just personal savings as a smoothing mechanism are clearly extreme characterization of individual behavior and of the economic environment faced by the consumers. Deaton and Paxson [1994] notice this and envision "the

[1] Even with precautionary saving, permanent shocks to labour income will typically be almost fully transmitted into consumption (see below).

construction and testing of market models under partial insurance", while Hayashi, Altonji and Kotlikoff [1996] call for future research to be "directed to estimating the extent of consumption insurance over and above self-insurance". In this paper we start from the premise of some, but not necessarily full, insurance and consider the importance of distinguishing between transitory and permanent shocks. We address the issue of whether partial consumption insurance is available to agents and estimate the degree of insurance over and above self-insurance through savings. Our research is related to other papers in the literature, particularly Hall and Mishkin [1982], Altonji, Martins and Siow [2002], Deaton and Paxson [1994], Moffitt and Gottschalk [1994], and Blundell and Preston [1998].[2]

Our data combine information from the Panel Study of Income Dynamics (PSID) and the Consumer Expenditure Survey (CEX) to document a number of key findings. We find a strong growth in permanent income shocks during the early 1980s. The variance of permanent shocks thereafter levels off. We find compelling evidence against full insurance for permanent income shocks but not for transitory income shocks, except for a low income subsample where transitory shocks seem, unsurprisingly, less insurable. Further there is evidence of some *partial insurance* of permanent income shocks, the degree of which varies across demographic groups. We find that consumption inequality − a topic that, with the notable exceptions of Cutler and Katz [1992] and Dynarski and Gruber [1997], has been studied much less extensively than wage inequality − follows closely the trends in permanent earnings inequality documented, among others, by Moffitt and Gottschalk [1994].[3] Our results point to durable expenditures being an important mechanism for

[2]Hall and Mishkin [1982] use panel data on food consumption and income from the PSID and consider the covariance restrictions imposed by the PIH. Altonji, Martins, and Siow [2002] improve on this by estimating a dynamic factor model of consumption, hours, wages, unemployment, and income, again using PSID data. Deaton and Paxson [1994] use repeated cross-section data from the US, the UK, and Thailand to test the implications that the PIH imposes on consumption inequality. Moffitt and Gottschalk [1994] use PSID panel on income to identify the variance of permanent and transitory income shocks. Blundell and Preston [1998] use the growth in consumption inequality over the 1980s in the U.K. to identify growth in permanent (uninsured) income inequality. Unlike Moffitt and Gottschalk [1995] −who use panel data on income but not consumption− they use data on both income and consumption but lack a panel dimension. Our use of panel data on income *and* consumption allows us to identify the variance of the income shocks as well as the degree of insurance of consumption with respect to the two types of shocks.

[3]The literature on consumption inequality is growing steadily. See, e.g., Attanasio, Battistin and Ichimura [2004],

smoothing non-durable consumption in the presence of income shocks especially for low income households. Finally we show that taxes and transfers provide an important insurance mechanism for permanent income shocks.

We use the term *partial insurance* to denote smoothing devices other than credit markets for borrowing and saving. There is scattered evidence on the role played by such devices on household consumption. Theoretical and empirical research have analyzed the role of extended family networks [Kotlikoff and Spivak, 1981; Attanasio and Rios-Rull, 2000], added worker effects [Stephens, 2002], the timing of durable purchases [Browning and Crossley, 2003], progressive income taxation [Mankiw and Kimball, 1992, Auerbach and Feenberg, 2001, and Kniesner and Ziliak, 2002], personal bankruptcy laws [Fay, Hurst and White, 2002], insurance within the firm [Guiso, Pistaferri and Schivardi, 2003], and the role of government public policy programs, such as unemployment insurance [Engen and Gruber, 2001], Medicaid [Gruber and Yelowitz, 1999], AFDC [Gruber, 2000], and food stamps [Blundell and Pistaferri, 2003]. While we do not take a precise stand on the mechanisms (other than savings) that are available to smooth idiosyncratic shocks to income, we emphasize that our evidence can be used to uncover whether some of these mechanisms are actually at work, how important they are quantitatively, and how they differ across households and over time. Our approach of examining the relationship between consumption and income inequality follows the suggestion of Deaton [1997] that "although it is possible to examine the mechanisms [providing partial insurance against income shocks], their multiplicity makes it attractive to look directly at the magnitude that is supposed to be smoothed, namely consumption".

The distinction between permanent and transitory shocks stressed in this paper is an important one, as we might expect to uncover less insurance for more persistent shocks. This point has been emphasized in the early work on the permanent income hypothesis and also in the recent wave of limited commitment models, which is one example where one might expect the relationship between income shocks and consumption to depend on the degree of persistence of income shocks.

Heathcote, Storesletten and Violante [2004], and Krueger and Perri [2003].

The literature on insurance under limited commitment [Kehoe and Levine, 2001, Alvarez and Jermann, 2000] explores the nature of income insurance schemes in economies where agents cannot be prevented from withdrawing participation if the loss from the accumulated income gains they are asked to forgo becomes greater than the gains from continuing participation. Such schemes, if feasible, allow individuals to keep some of the positive shocks to their income and therefore offer only partial income insurance. The proportion of income shocks which is insured will vary −among other things− with the variance of the underlying shocks. As the variance increases the value of future participation increases, alleviating the participation constraint.[4] This is particularly relevant in the US and the UK, where quantitatively large changes in the structure of relative prices (most notably, wages) have occurred over the last three decades, both within and between groups. The results in Alvarez and Jermann also demonstrate that if income shocks are persistent enough and agents are infinitely lived, then participation constraints become so severe that no insurance scheme is feasible. This suggests that the degree of insurance should be allowed to differ between transitory and permanent shocks and should also be allowed to change over time and across different groups.

Uncovering the degree of partial insurance is likely to matter for a number of reasons. First, it may help to understand the characteristics of the economic environment faced by the agents. This may prove crucial when evaluating the performance of macroeconomic models, especially those that explicitly account for agents' heterogeneity. Moreover, it is important to understand to what extent changes in social insurance systems affect smoothing abilities, and the consequences of this for private saving behavior. This is important as far as the efficient design and evaluation of social insurance policy is concerned. Finally, the presence of mechanisms that allow households to smooth idiosyncratic shocks has a bearing on aggregation results [see Blundell and Stoker, 2004].

A study of this kind requires in principle good quality longitudinal data on household consumption and income. It is well known that the PSID contains longitudinal income data but the

[4] Krueger and Perri [2003] investigate insurance of transitory shocks through analytic solution of simple models and simulation of more complex cases and demonstrate the possibility that consumption variance can actually fall with an increase in the variance of income shocks.

information on consumption is scanty (limited to food and few more items). Our strategy is to impute consumption to all PSID households combining PSID data with consumption data from repeated CEX cross-sections. Previous studies [Skinner, 1987] impute non-durable consumption data in the PSID using CEX regressions of non durable consumption on consumption items (food, housing, utilities) and demographics available in both the PSID and the CEX. Although related, our approach starts from a standard demand function for food (a consumption item available in both surveys); we make this depend on prices, total non durable expenditure, and a host of demographic and socio-economic characteristics of the household. Food expenditure and total expenditure are modeled as jointly endogenous. Under monotonicity (normality) of food demands these functions can be inverted to obtain a measure of non durable consumption in the PSID. In a companion paper [Blundell, Pistaferri and Preston, 2004] we review the conditions that make this procedure reliable and show that it is able to reproduce remarkably well the trends in the consumption distribution.

The paper continues with an illustration of the model we estimate and of the identification strategy we use (Section 2). In Section 3 we discuss data issues and the imputation procedure. Section 4 presents the empirical results and Section 5 concludes.

2 Income and Consumption Dynamics

2.1 The income process

The unit of analysis is a household, comprising a couple and, if present, their children. Our sample selection focuses on income risk and we do not model divorce, widowhood, and other household breaking-up factors. We recognize that these may be important omissions that limit the interpretation of our study. However, by focusing on stable households and the interaction of consumption and income we are able to develop a complete identification strategy.[5] We also select

[5] Whether stable families have access to more or less insurance than non-stable families is an issue that cannot be settled in principle. On the one hand, stable families have often more incomes and assets and therefore are less likely to be eligible for social insurance, which is typically means-tested. On the other hand, they can plausibly be more successful in securing access to credit, family networks and other informal insurance devices, over and above self-insurance through saving.

households during the working life of the husband.

We assume that the sole relevant source of uncertainty faced by the consumer is income (defined as the sum of labor income and transfers, such as welfare payments). We also assume that labor is supplied inelastically and make the assumption of separability in preferences between consumption and leisure. This means all insurance provided through, say, an added worker effect, will pass through disposable income. Similarly, it is possible that the wage component of family income may have already been smoothed out relative to productivity by implicit agreements within the firm. If this insurance is present, it will be reflected in the variability of income.

The income process for each household i we consider is:

$$\log Y_{i,a,t} = Z'_{i,a,t}\varphi_t + H_{i,a,t} + v_{i,a,t} \tag{1}$$

where a and t index age and time respectively, Y is real income, and Z is a set of income characteristics observable and anticipated by consumers. (Note that we allow the effect of such characteristics to shift with calendar time.) Equation (1) decomposes the remainder of income into a permanent component $H_{i,a,t}$ and a transitory or mean-reverting component, $v_{i,a,t}$. By writing $Y_{i,a,t}$ rather than $Y_{i,t}$ we emphasize the importance of cohort effects in the evolution of earnings over the life-cycle. In keeping with this remark, we also study consumption decisions of different cohorts separately.

For consistency with previous empirical studies [MaCurdy, 1982; Abowd and Card, 1989; Moffitt and Gottschalk, 1994; Meghir and Pistaferri, 2004], we assume that the permanent component $H_{i,a,t}$ follows a martingale process of the form:

$$H_{i,a,t} = H_{i,a-1,t-1} + \zeta_{i,a,t} \tag{2}$$

where $\zeta_{i,a,t}$ is serially uncorrelated, and the transitory component $v_{i,a,t}$ follows an MA(q) process, where the order q is to be established empirically:

$$v_{i,a,t} = \sum_{j=0}^{q} \theta_j \varepsilon_{i,a-j,t-j}$$

with $\theta_0 \equiv 1$. It follows that (unexplained) income growth is

$$\Delta y_{i,a,t} = \zeta_{i,a,t} + \Delta v_{i,a,t} \tag{3}$$

where $y_{i,a,t} = \log Y_{i,a,t} - Z'_{i,a,t}\varphi_t$ denotes the log of real income net of predictable individual components.

2.2 Insurance and the Transmission of Income Shocks to Consumption

2.2.1 Self Insurance

Consider the optimization problem faced by household i. The objective is to:

$$\max_C E_{a,t} \sum_{j=0}^{T-a} u\left(C_{i,a+j,t+j}\right) e^{Z'_{i,a+j,t+j}\vartheta_{t+j}} \tag{4}$$

where $Z'_{i,a+j,t+j}\vartheta_{t+j}$ incorporates taste shifters and discount rate heterogeneity. Maximization of (4) is subject to the budget constraint

$$A_{i,a+j+1,t+j+1} = \left(1 + r_{t+j}\right)\left(A_{i,a+j,t+j} + Y_{i,a+j,t+j} - C_{i,a+j,t+j}\right) \tag{5}$$

$$A_{i,T,t+T-a} = 0 \tag{6}$$

with $A_{i,a,t}$ given. We set the retirement age after which income falls to zero at L, assumed known and certain, and the end of the life-cycle at age T. We assume that there is no interest rate uncertainty or uncertainty about the date of death. If preferences are of the CRRA form ($u\left(C\right) = \frac{C^{\beta}-1}{\beta}$) and credit markets are perfect, then optimal consumption choices can be described by an approximate consumption growth equation, derived in Appendix A.1, which provides a mapping from the income shocks $\zeta_{i,a,t}$ and $\varepsilon_{i,a,t}$ to the optimal consumption growth, given by

$$\Delta c_{i,a,t} \cong \pi_{i,a,t}\zeta_{i,a,t} + \pi_{i,a,t}\gamma_{t,L}\varepsilon_{i,a,t} + \xi_{i,a,t} \tag{7}$$

where $\Delta c_{i,a,t} = \Delta \log C_{i,a,t} - \Delta Z'_{i,a,t}\vartheta_t - \Gamma_{b,t}$ is the log of real consumption net of its predictable components. Appendix A.1 shows that the term $\Gamma_{b,t}$ is the slope of the consumption path for the individual's year-of-birth cohort (which we index with b), while $\xi_{i,a,t}$ is a random term that can be interpreted as the individual deviation from the cohort-specific consumption gradient.[6] The coefficients on the income shocks are determined by $\pi_{i,a,t}$, which is the share of future labor income in the present value of lifetime wealth, and $\gamma_{t,L}$, which is an age-increasing known weight.[7]

Interpretation of the impact of income shocks on consumption growth is straightforward. For individuals a long time from the end of their life with the value of current financial assets small relative to remaining future labor income, $\pi_{i,a,t} \simeq 1$, and permanent shocks pass through more or less completely into consumption whereas transitory shocks are (almost) completely insured against through saving. This is the main insight of the textbook permanent income hypothesis [Deaton, 1992]. Precautionary saving can provide effective self-insurance against permanent shocks only if the stock of assets built up is large relative to future labor income, which is to say $\pi_{i,a,t}$ is appreciably smaller than unity, in which case there will also be some smoothing of permanent shocks through self insurance (see also Carroll, 2001, for numerical simulations).

2.2.2 Additional Insurance

While precautionary saving might allow some insurance of permanent shocks if assets are large enough relative to future labor income (i.e. $\pi_{i,a,t} < 1$) other interpersonal insurance mechanisms might also underlie this. We now consider the possibility of additional insurance and suppose there are mechanisms (that we do not model explicitly here but were discussed in the Introduction) that allow insurance of a fraction $(1 - \phi_{b,t})$ and $(1 - \psi_{b,t})$ of permanent and transitory shocks, respectively. We might expect $\phi_{b,t}$ to be close to unity and $\psi_{b,t}$ close to zero.[8]

In this case consumption growth can be written as:

$$\Delta c_{i,a,t} \cong \phi_{b,t}\zeta_{i,a,t} + \psi_{b,t}\varepsilon_{i,a,t} + \xi_{i,a,t} \tag{8}$$

The economic interpretation of the partial insurance parameter is such that it nests the two polar

[6] Innovations to the conditional variance of consumption growth (precautionary savings) are captured by $\Gamma_{b,t}$.

[7] See Appendix A.1. Results from a simulation of a stochastic economy presented in Blundell, Low and Preston (2004) show that this approximation can be used to accurately detect changes in the time series pattern of permanent and transitory variances to income shocks. These results are available on request (by email to: i.preston@ucl.ac.uk).

[8] If there are no interpersonal mechanisms or transfers of any sort, then $\phi_{b,t} = \psi_{b,t} = \pi_{i,a,t} = 1$.

cases of full insurance of income shocks ($\phi_{b,t} = \psi_{b,t} = 0$), as contemplated by the complete markets hypothesis, and no insurance ($\phi_{b,t} = \psi_{b,t} = 1$), as well as the intermediate case $\phi_{b,t} = \psi_{b,t}/\gamma_{t,L} = \pi_{i,a,t}$ predicted by the PIH with self-insurance through savings. A value $0 < \phi_{b,t} < 1$ ($0 < \psi_{b,t} < 1$) is consistent with partial insurance with respect to permanent (transitory) shocks. The lower the coefficient, the higher the degree of insurance.

2.2.3 Advance Information

In the analysis presented so far we have assumed that in the innovation process for income (3) the random variables $\zeta_{i,a,t}$ and $\varepsilon_{i,a,t}$ represent the arrival of new information to the agent i of age a in period t. If parts of these random terms were known in advance to the agent then the consumption model would argue that they should already be incorporated into consumption plans and would not directly effect consumption growth (8). Suppose, for example, that only a proportion κ of the permanent shock was unknown to the consumer. Then the consumption growth relationship (8) would become

$$\Delta c_{i,a,t} \cong \widetilde{\phi}_{b,t}\ \kappa\ \zeta_{i,a,t} + \psi_{b,t}\varepsilon_{i,a,t} + \xi_{i,a,t}. \tag{9}$$

where $\widetilde{\phi}_{b,t}$ is the "true" insurance parameter. In this case $\phi_{b,t}$ would be underestimated by the information factor κ.

The econometrician will treat $\zeta_{i,a,t}$ as the permanent shock. Whereas the individual may have already adapted to this change. Consequently, although transmission of income inequality to consumption inequality is correctly identified, the estimated $\phi_{b,t}$ has to be interpreted as reflecting a combination of insurance and information. In the absence of outside information (such as, say, subjective expectations), these two components cannot be separately identified. The issue is discussed further in Section 4 where we interpret our empirical results.

When we allow for partial insurance or advance information, we are unable to separately identify how precautionary saving (through $\pi_{i,a,t}$) and either partial insurance over and above saving or foresight smooth the impact of shocks on consumption. However, this will be practically of little importance. We will be identifying a parameter that combines self-insurance, partial insurance, foresight and perhaps even the crowding out effect of public insurance on private insurance. In other words, our generalised partial insurance parameters will still pin down the degree of transmission of income shocks into consumption, which is our primary objective.

2.3 Evolution of Income and Consumption Variances

We assume that $\zeta_{i,a,t}$, $v_{i,a,t}$ and $\xi_{i,a,t}$ are mutually uncorrelated processes. Equation (3) can be used to derive the following covariance restrictions in panel data

$$\operatorname{cov}(\Delta y_{a,t}, \Delta y_{a+s,t+s}) = \begin{cases} \operatorname{var}(\zeta_{a,t}) + \operatorname{var}(\Delta v_{a,t}) & \text{for } s = 0 \\ \operatorname{cov}(\Delta v_{a,t}, \Delta v_{a+s,t+s}) & \text{for } s \neq 0 \end{cases} \tag{10}$$

where var (.) and cov (., .) denote cross-sectional variances and covariances, respectively (the index i is consequently omitted). These moments can be computed for the whole sample or for individuals belonging to a homogeneous group (i.e., born in the same year, with the same level of schooling, etc.). The covariance term $\operatorname{cov}(\Delta v_{a,t}, \Delta v_{a+s,t+s})$ depends on the serial correlation properties of v. If v is an $MA(q)$ serially correlated process, then $\operatorname{cov}(\Delta v_{a,t}, \Delta v_{a+s,t+s})$ is zero whenever $|s| > q+1$. Note also that if v is serially uncorrelated ($v_{i,a,t} = \varepsilon_{i,a,t}$), then $\operatorname{var}(\Delta v_{a,t}) = \operatorname{var}(\varepsilon_{a,t}) + \operatorname{var}(\varepsilon_{a-1,t-1})$. See also Moffitt and Gottschalk [1994]. Identification of the serial correlation coefficients does not hinge on the order of the process q. Allowing for an MA(q) process, for example, adds $q-1$ extra parameter (the $q-1$ MA coefficients) but also $q-1$ extra moments, so that identification is unaffected.

The panel data restrictions on consumption growth from (8) are as follows:

$$\operatorname{cov}(\Delta c_{a,t}, \Delta c_{a+s,t+s}) = \phi_{b,t}^2 \operatorname{var}(\zeta_{a,t}) + \psi_{b,t}^2 \operatorname{var}(\varepsilon_{a,t}) + \operatorname{var}(\xi_{a,t}) \tag{11}$$

for $s = 0$ and zero otherwise (due to the consumption martingale assumption).

Finally, the covariance between income growth and consumption growth at various lags is:

$$\operatorname{cov}(\Delta c_{a,t}, \Delta y_{a+s,t+s}) = \begin{cases} \phi_{b,t} \operatorname{var}(\zeta_{a,t}) + \psi_{b,t} \operatorname{var}(\varepsilon_{a,t}) \\ \psi_{b,t} \operatorname{cov}(\varepsilon_{a,t}, \Delta v_{a+s,t+s}) \end{cases} \tag{12}$$

for $s = 0$, and $s > 0$ respectively. If v is an $MA(q)$ serially correlated process, then $\text{cov}(\Delta c_{a,t}, \Delta y_{a+s,t+s})$ is zero whenever $|s| > q+1$. Thus, if v is serially uncorrelated ($v_{i,a,t} = \varepsilon_{i,a,t}$), then $\text{cov}(\Delta c_{a,t}, \Delta y_{a+s,t+s}) = -\psi_{b,t}\text{var}(\varepsilon_{a,t})$ for $s = 1$ and 0 otherwise.

Note finally that it is likely that measurement error will contaminate the observed income and consumption data. Assume that both consumption and income are measured with multiplicative independent errors, e.g.,

$$y^*_{i,a,t} = y_{i,a,t} + u^y_{i,a,t} \tag{13}$$

and

$$c^*_{i,a,t} = c_{i,a,t} + u^c_{i,a,t} \tag{14}$$

where x^* denote a measured variable, x its true, unobservable value, and u the measurement error. In Appendix A.2 we show that the partial insurance parameter $\phi_{b,t}$ remains identified under measurement error, while only a lower bound for $\psi_{b,t}$ is identifiable. A corollary of this is that the variance of measurement error in consumption can be identified (the theory suggests that consumption should be a martingale with drift, so any serial correlation in consumption growth can only be attributed to noise), but the variance of the measurement error in income can still not be identified separately from the variance of the transitory shock.[9] The goal of the empirical analysis is to estimate features of the distribution of income shocks (variances of permanent and transitory shocks and the extent of serial correlation in the latter) and consumption growth (particularly the partial insurance parameters) using joint panel data on income and consumption growth on which the theoretical restrictions (10)-(12) have been imposed.

In the context of identifying sources of variation in household income and consumption, it is worth stressing that in addition to identifying the partial insurance parameters, the availability of panel data presents several advantages over a repeated cross-sections analysis. With repeated cross sections the variances and covariances of differences in income and consumption cannot be observed,

[9] Thus the variance of measurement error in consumption is identified by $-\text{cov}(\Delta c_{a,t}, \Delta c_{a+1,t+1})$.

though it is possible to make assumptions under which variances of shocks can be identified from differences in variances and covariances of their levels. For example, under the assumption that shocks are cross-sectionally orthogonal to past consumption and income and that transitory shocks are serially uncorrelated, Blundell and Preston [1998] use repeated cross-section moments to separate the growth in the variance of transitory shocks to log income from the variance of permanent shocks (see also Deaton and Paxson [1994]). This orthogonality assumption will be violated if, say, knowledge of one's position in the income (or consumption) distribution conveys information about the distribution of future shocks to income. In panel data, identification does not require making such assumption and can allow for serial correlation in transitory shocks as well as measurement error in consumption and income data (see below).

With panel data the identification of the variances of shocks to income requires only panel data on income, not consumption. In the simple case of serially uncorrelated transitory shock, for example:[10]

$$\operatorname{var}(\zeta_{a,t}) = \operatorname{cov}(\Delta y_{a,t}, \Delta y_{a-1,t-1} + \Delta y_{a,t} + \Delta y_{a+1,t+1}) \tag{15}$$

$$\operatorname{var}(\varepsilon_{a,t}) = -\operatorname{cov}(\Delta y_{a,t}, \Delta y_{a+1,t+1}) \tag{16}$$

Using panel data on both consumption and income improves efficiency of these estimates because it provides extra moments for identification. We will show that the two sets of estimates are basically the same. The joint use of consumption and income data allows identification of the insurance parameters that would not be identifiable with income or consumption data used in isolation. In turn, knowledge of the extent of insurance is informative about the welfare effects of shifts in the income distribution.

[10] See Meghir and Pistaferri [2004] for a generalization to serially correlated transitory shocks and measurement error in income.

3 The data

Our empirical analysis combines microeconomic data from two sources: the 1978-1992 PSID and the 1980-1992 CEX. We describe their main features and our sample selection procedures in turn.

3.1 The PSID

Since the PSID has been widely used for microeconometric research, we shall only sketch the description of its structure in this section.[11]

The PSID started in 1968 collecting information on a sample of roughly 5,000 households. Of these, about 3,000 were representative of the US population as a whole (the core sample), and about 2,000 were low-income families (the Census Bureau's Survey of Economic Opportunities, or SEO sample). Thereafter, both the original families and their split-offs (children of the original family forming a family of their own) have been followed.

The PSID includes a variety of socio-economic characteristics of the household, including education, food spending, and income of household members. Questions referring to income are retrospective; thus, those asked in 1993, say, refer to the 1992 calendar year. In contrast, the timing of the survey questions on food expenditure is much less clear [see Hall and Mishkin, 1982, and Altonji and Siow, 1987, for two alternative views]. Typically, the PSID asks how much is spent on food in an average week. Since interviews are usually conducted around March, it has been argued that people report their food expenditure for an average week around that period, rather than for the previous calendar year as is the case for family income. We assume that food expenditure reported in survey year t refers to the previous calendar year, but check the effect of alternative assumptions.

Households in the PSID report their taxable family income (which includes transfers and financial income). The measure of income used in the baseline analysis below excludes income from financial assets, subtracts taxes and deflates the corresponding value by the CPI. We obtain an

[11] See Hill [1992] for more details about the PSID.

after-tax measure of income subtracting federal taxes paid. Before 1991, these are computed by PSID researchers and added into the data set using information on filing status, adjusted gross income, whether the respondent itemizes or takes the standard deduction, and other household characteristics that make them qualify for extra deductions, exemptions, and tax credits. Federal taxes are not computed in 1992 and 1993. We impute taxes for the last two years using regression analysis for the years where taxes are available (results not reported but available on request).

Education level is computed using the PSID variable "grades of school finished". Individuals who changed their education level during the sample period are allocated to the highest grade achieved. We consider two education groups: with and without college education (corresponding to 13 grades or more and 12 grades or less, respectively).

Since CEX data are available on a consistent basis since 1980, we construct an unbalanced PSID panel using data from 1978 to 1992 (the first two years are retained for initial conditions purposes). Due to attrition, changes in family composition, and various other reasons, household heads in the 1978-1992 PSID may be present from a minimum of one year to a maximum of fifteen years. We thus create unbalanced panel data sets of various length. The longest panel includes individuals present from 1978 to 1992; the shortest, individuals present for two consecutive years only (1978-79, 1979-80, up to 1991-92).

The objective of our sample selection is to focus on a sample of continuously married couples headed by a male (with or without children). The step-by-step selection of our PSID sample is illustrated in Table I. We eliminate households facing some dramatic family composition change over the sample period. In particular, we keep only those with no change, and those experiencing changes in members other than the head or the wife (children leaving parental home, say). We next eliminate households headed by a female, those with missing report on education and region,[12] and those with topcoded income. We keep continuously married couples and drop some income

[12] When possible, we impute values for education and region of residence using adjacent records on these variables.

outliers.[13] We then drop those born before 1920 or after 1959.

As noted above, the initial 1967 PSID contains two groups of households. The first is representative of the US population (61 percent of the original sample); the second is a supplementary low income subsample (also known as SEO subsample, representing 39 percent of the original 1967 sample). For the most part we exclude SEO households and their split-offs. However, we do consider the robustness of our results in the low income SEO subsample.

Finally, we drop those aged less than 30 or more than 65. This is to avoid problems related to changes in family composition and education, in the first case, and retirement, in the second. The final sample used in the minimum distance exercise below is composed of 17,788 observations and 1,788 households.

We use information on age and the survey year to allocate individuals in our sample to four cohorts defined on the basis of the year of birth of the household head: born in the 1920s, 1930s, 1940s, and 1950s. Years where cell size is less than 100 are discarded.[14]

3.2 The CEX

The Consumer Expenditure Survey provides a continuous and comprehensive flow of data on the buying habits of American consumers. The data are collected by the Bureau of Labor Statistics and used primarily for revising the CPI.[15] The definition of the head of the household in the CEX is the person or one of the persons who owns or rents the unit; this definition is slightly different from the one adopted in the PSID, where the head is always the husband in a couple. We make the two definitions compatible.

The CEX is based on two components, the Diary survey and the Interview survey. The Diary sample interviews households for two consecutive weeks, and it is designed to obtain detailed expen-

[13] An income outlier is defined as a household with an income growth above 500 percent, below -80 percent, or with a level of income below \$100 a year or below the amount spent on food.

[14] Median (average) cell sizes are 249 (219), 245 (246), 413 (407), and 398 (363), respectively for those born in the 1920s, 1930s, 1940s, and 1950s.

[15] A description of the survey, including more details on sample design, interview procedures, etc., may be found in "Chapter 16: Consumer Expenditures and Income", from the BLS Handbook of Methods.

ditures data on small and frequently purchased items, such as food, personal care, and household supplies. The Interview sample follows survey households for a maximum of 5 quarters, although only inventory and basic sample data are collected in the first quarter. The data base covers about 95% of all expenditure, with the exclusion of expenditures for housekeeping supplies, personal care products, and non-prescription drugs. Following most previous research, our analysis below uses only the Interview sample.[16]

As the PSID, the CEX collects information on a variety of socio-demographic variables, including income and consumer expenditure. Expenditure is reported in each quarter and refers to the previous quarter; income is reported in the second and fifth interview (with some exceptions), and refers to the previous twelve months. For consistency with the timing of consumption, fifth-quarter income data are used.

We select a CEX sample that can be made comparable, to the extent that this is possible, to the PSID sample. Our initial 1980-1998 CEX sample includes 1,249,329 monthly observations, corresponding to 141,289 households. We drop those with missing record on food and/or zero total nondurable expenditure, and those who completed less than 12 month interviews. This is to obtain a sample where a measure of annual consumption can be obtained. A problem is that many households report their consumption for overlapping years, i.e. there are people interviewed partly in year t and partly in year $t+1$. Pragmatically, we assume that if the household is interviewed for at least 6 months at $t+1$, then the reference year is $t+1$, and it is t otherwise. Prices are adjusted accordingly. We then sum food at home, food away from home and other nondurable expenditure over the 12 interview months. This gives annual expenditures. For consistency with the timing of the PSID data, we drop households interviewed after 1992. We also drop those with zero before-tax income, those with missing region or education records, single households and those with changes in family composition. Finally, we eliminate households where the head is born before 1920 or

[16] There is some evidence that trends in consumption inequality measured in the two CEX surveys have diverged in the 1990s [Attanasio, Battistin and Ichimura, 2004]. While research on the reasons for this divergence is clearly warranted, our analysis, which uses data up to 1992, will only be marginally affected.

after 1959, those aged less than 30 or more than 65, and those with outlier income (defined as a level of income below the amount spent on food) or incomplete income responses. Our final sample contains 15,137 households. Table II details the sample selection process in the CEX.

The definition of total non durable consumption is the same as in Attanasio and Weber [1995]. It is the sum of food (defined as the sum of food at home and food away from home), alcohol, tobacco, and expenditure on other nondurable goods, such as services, heating fuel, public and private transports (including gasoline), personal care, and semidurables, defined as clothing and footwear. This definition excludes expenditure on various durables, housing (furniture, appliances, etc.), health, and education. In our empirical results we assess the sensitivity of our results to the inclusion of durables and other non-durable items.[17]

3.3 Comparing and combining the two data sets

How similar are the two data sets in terms of average demographic and socio-economic characteristics? Mean comparisons are reported in Table III for selected years: 1980, 1983, 1986, 1989, and 1992. The PSID respondents are slightly younger than their CEX counterparts; there is, however, little difference in terms of family size and composition. The percentage of whites is slightly higher in the PSID. The distribution of the sample by schooling levels is quite similar, while the PSID tends to under-represent the proportion of people living in the West. Both male and female participation rates in the PSID are comparable to those in the CEX. Due to slight differences in the definition of family income, PSID figures are higher than those in the CEX. It is possible that the definition of family income in the PSID is more comprehensive than that in the CEX, so resulting in the underestimation of income in the CEX that appears in the Table. Total food expenditure (the sum of food at home and food away from home) is fairly similar in the two data sets. Blundell, Pistaferri and Preston [2004] provides a detailed comparison of the components of the total food consumption series.

[17] We also tried with a definition of nondurable consumption that includes services from durables (housing and vehicles). We thanks David Johnson at BLS for providing data on the latter.

In deriving the theoretical restrictions in Section 2, we have assumed that a researcher has access to panel data on household income and total non-durable consumption. However, this is a very strong data requirement. In the US, panel data typically lack household data on total non-durable consumption; and those surveys, such as the CEX, that contains good quality data on consumption, lack a panel feature. We may however combine the two data sets to impute non durable consumption to PSID households.[18]

The PSID collects data on few consumption items, mainly food at home and food away from home. Moreover, food data are not available in 1987 and 1988. Our strategy is to write a demand equation for food as a function of prices, demographics, and total non-durable expenditure. We then use the inverse demand to obtain an imputed measure of total non-durable consumption. This inversion operation requires consistent estimation of the parameters of the demand function for food and monotonicity of the underlying demand function.

The technical details of the imputation procedure and a sequence of robustness tests are provided in Blundell, Pistaferri and Preston [2004]. Briefly, we pool all the CEX data from 1980 to 1992, and write the following demand equation for food

$$f_{i,a,t} = W'_{i,a,t}\mu + \beta\left(D_{i,a,t}\right)c_{i,a,t} + e_{i,a,t} \tag{17}$$

where f is the log of food expenditure (which is available in both surveys), W contains prices and a set of demographic variables (also available in both data sets), c is the (endogenous) log of total non-durable expenditure (available only in the CEX), and e captures unobserved heterogeneity in the demand for food and measurement error in food expenditure. We allow for the elasticity $\beta(.)$ to vary with time and with observable household characteristics. The estimation results for our specification of (17) are reported in Table IV. To account for measurement error and general

[18] Previous studies [Skinner, 1987] impute non-durable consumption data in the PSID using CEX regressions of non durable consumption on consumption items available in both data sets. The only consumption items that are available in the PSID on a consistent basis are food expenditure and rents (in the early years of the survey many more items were available, such as utilities, alcohol, tobacco, child care, transport costs to work, and car insurance, but their collection was discontinued mostly after 1972). Given that the majority of households own their home, the rent variable must be imputed. If one is unwilling to use this variable, the Skinner procedure and the one we suggest here (apart from our emphasis on controlling for prices and demographics) are very similar.

endogeneity of total expenditure we instrument the latter with the average (by cohort, year, and education) of the hourly wage of the husband and the average (also by cohort, year, and education) of the hourly wage of the wife. The budget elasticity is 0.88 (0.81 in the OLS case). The price elasticity is -0.96. We test the overidentifying restrictions and fail to reject the null hypothesis (p-value of 56 percent). We also report statistics for judging the power of excluded instruments. They are all acceptable. Generally the demographics have the expected sign.

For the purposes of this study a good inversion procedure should have the property that the variance of (imputed) consumption in the PSID should exceed the variance of consumption in the CEX by an additive factor (the variance of the error term of the demand equation scaled by the square of the expenditure elasticity). If this factor is constant over time the *trends* in the two variances should be identical. We refer the interested reader to Blundell, Pistaferri and Preston [2004] for more details. Figure 1 shows that the variances line up extremely well. The range of variation of the variance of PSID consumption is on the left-hand side; that of the variance of CEX consumption, on the right hand side. Trends in the variance of consumption are remarkably similar in the two data sets. In fact the reader can check that the variance of imputed PSID consumption is just an upward-translated version (by about 0.05 units) of the variance of CEX consumption. Both series suggest that between 1980 and 1986 the variance of log consumption (a standard measure of consumption inequality) grows quite substantially. Afterwards, both graphs are flat. In Blundell, Pistaferri and Preston [2004] we show that this result is robust to variation in equivalence scale; we also show that our imputation procedure is capable of replicating quite well the trends in mean spending as long as account is made for differences in the mean of the input variable (food spending) in the two data sets.

The evidence discussed in this section thus provides confidence in our use of imputed data to estimate the parameters of interest discussed in Section 2. We now turn to the results of our empirical analysis.

4 The results

We first discuss the characterization of the variance-covariance structure of consumption and income in the PSID. We then evaluate the relative size and trends in the variance of permanent and transitory shocks to income and estimate the degree of insurance to these shocks for different sub-groups of the population.

4.1 Autocovariance Estimates of Consumption and Income: Longitudinal Evidence from the Matched PSID

The PSID data set contains longitudinal records on income and imputed consumption. We remove the effect of deterministic effects on log income and (imputed) log consumption by separate regressions of these variables on year and year of birth dummies, and on a set of observable family characteristics (dummies for education, race, family size, number of children, region, employment status, residence in a large city, outside dependent, and presence of income recipients other than husband and wife). We allow for the effect of these characteristics to vary with calendar time. These variables are assumed to reflect deterministic growth in consumption and income (e.g., information). We then work with the residuals of these regressions, $c_{i,a,t}$ and $y_{i,a,t}$.

To pave the way to the formal analysis of partial insurance, Table V reports unrestricted minimum distance estimates of several moments of the income process for the whole sample: the variance of unexplained income growth, $\text{var}\left(\Delta y_{a,t}\right)$, the first-order autocovariances $\left(\text{cov}\left(\Delta y_{a+1,t+1}, \Delta y_{a,t}\right)\right)$, and the second-order autocovariances $\left(\text{cov}\left(\Delta y_{a+2,t+2}, \Delta y_{a,t}\right)\right)$. Estimates are reported for each year. Table VI repeats the exercise for our measure of consumption. Finally, Table VII reports minimum distance estimates of contemporaneous and lagged consumption-income covariances. Some of the moments are missing because, as said above, consumption data were not collected in the PSID in 1987-88.

Looking through Table V, one can notice the strong increase in the variance of income growth, rising by more than 30% by 1986. Also notice the strong blip in the final year (in 1992 the PSID converted the questionnaire to electronic form and imputations of income done by machine). The absolute value of the first-order autocovariance also increases through to 1986 and then is stable or even declines. Second- and higher order autocovariances (which, from equation (10), are informative about the presence of serial correlation in the transitory income component) are small and only in few cases statistically significant. At least at face value, this evidence seems to tally quite well with a canonical MA(1) process in growth, as implied by a traditional income process given by the sum of a martingale permanent component and a serially uncorrelated transitory component. Since evidence on second-order autocovariances is mixed, however, in estimation we allow for MA(1) serial correlation in the transitory component ($v_{i,a,t} = \varepsilon_{i,a,t} + \theta\varepsilon_{i,a-1,t-1}$).

While income moments are informative about shifts in the income distribution (and on the temporary or persistent nature of such shifts), they cannot be used to make conclusive inference about shifts in the consumption distribution. For this purpose, one needs to complement the analysis of income moments with that of consumption moments and of the joint income-consumption moments. This is done in Tables VI and VII. Table VI shows that the variance of imputed consumption growth also increases quite strongly in the early 1980s, peaks in 1985 and then it is essentially flat afterwards. Note the high value of the level of the variance which is clearly the result of our imputation procedure. The variance of consumption growth captures in fact the genuine association with shocks to income, but also the contribution of slope heterogeneity and measurement error.[19] The absolute value of the first-order autocovariance of consumption growth should be a good estimate of the variance of the imputation error. This is in fact quite high and approximately stable over time. Second-order consumption growth autocovariances are mostly statistically insignificant and economically small.

Table VII looks at the association, at various lags, of unexplained income and consumption growth. The contemporaneous covariance should be informative about the effect of income shocks

[19] To a first approximation, the variance of consumption growth that is not contaminated by error can be obtained by subtracting twice the (absolute value of) first order autocovariance $\text{cov}(\Delta c_{t+1}, \Delta c_t)$ from the variance $\text{var}(\Delta c_t)$.

on consumption growth if measurement errors in consumption are orthogonal to measurement errors in income. This covariance increases in the early 1980s and then is flat or even declining afterwards.

>From (14.6), the covariance between current consumption growth and future income growth $\text{cov}(\Delta c_{a,t}, \Delta y_{a+1,t+1})$ should reflect the extent of insurance with respect to transitory shocks (i.e., $\text{cov}(\Delta c_{a,t}, \Delta y_{a+1,t+1}) = 0$ if there is full insurance of transitory shocks). We note that in the pure self-insurance case with infinite horizon and MA(1) transitory component, the impact of transitory shocks on consumption growth is given by the annuity value $\frac{r(1+r-\theta)}{(1+r)^2}$. With a small interest rate, this will be indistinguishable from zero, at least statistically. In fact, this covariance is hardly statistically significant and economically close to zero. As we shall see, the formal analysis below will confirm this. We should note, however, that for the low income sample examined further in the empirical results below we do find some sensitivity to transitory shocks.

The covariance between current consumption growth and past income growth $\text{cov}(\Delta c_{a+1,t+1}, \Delta y_{a,t})$ plays no role in the PIH model with perfect capital markets, but may be important in alternative models where liquidity constraints are present (a standard excess sensitivity argument, see Flavin [1981]). The estimates of this covariance in Table VII are close to zero.

To sum up, there is weak evidence that transitory shocks impact consumption growth or that liquidity constraints are empirically important in this sample. In the sensitivity results reported below we note that there is more evidence of responsiveness to transitory shocks for the low income poverty sample of the PSID. We now turn to more formal minimum distance estimation, where we impose the theoretical restrictions outlined in Section 2.3 on the unrestricted income and consumption moments of Table V, VI, and VII.

4.2 Partial Insurance

Our estimates are based on a generalization of moments (10)-(12). In particular, to account for our imputation procedure, we assume that consumption is measured with error. We estimate the variance of the measurement error ($\sigma^2_{u^c}$) assuming that it is i.i.d. We also consider an MA(1) process for the transitory error component of income ($v_{i,a,t} = \varepsilon_{i,a,t} + \theta\varepsilon_{i,a-1,t-1}$), and estimate the MA(1) parameter θ. Finally, we allow for i.i.d. unobserved heterogeneity in the individual consumption gradient, and estimate its variance (σ^2_ξ). We present the results of three specifications: one for the whole sample (the "baseline" specification), one where parameters are estimated separately by education (college vs. no college), and one where parameters are estimated separately by cohort (born 1930s vs. born 1940s).[20] We also allow for some time non-stationarity. In particular, in all specifications we let the variances of the permanent and the transitory shock, σ^2_ζ and σ^2_ε, respectively, vary with calendar time. As for the partial insurance coefficients for the permanent shock (ϕ) and for the transitory shock (ψ), we assume that they take on two different values, before and after 1985. This is consistent with the the evidence in Figure 1, which divides the sample period in a period of rapid growth in the variance (up until 1985), and one of relative stability afterwards. We test the null that the extent of insurance does not change over time, and with almost no exceptions we fail to reject the null. Tables VIII, IX, and X will thus present the results of a simple model in which the insurance parameters are constant over time. We comment on the time variability of the insurance parameters where appropriate and present the results of the test in the tables.

The parameters are estimated by diagonally weighted minimum distance (DWMD). This estimation method is a simple generalization of equally minimum distance (EWMD). Unlike EWMD, it allows for heteroskedasticity. Moreover, it avoids the pitfalls of optimal minimum distance (OMD) remarked by Altonji and Segal [1996], which are primarily related to the terms outside the main diagonal of the optimal weighting matrix. Technical details are in Appendix A.3.[21]

The first column of Table VIII shows the results for the whole sample. The estimated variances

[20] Results for the younger cohort (born in the 1950s) and the older cohort (born in the 1920s) are less reliable because these cohorts are not observed for the whole sample period. We thus omit them.

[21] If we use EWMD we obtain extremely downward biased estimates of $\text{var}(\zeta_{a,t})$ and extremely upward biased estimates of $\text{var}(\varepsilon_{a,t})$ (compared to those we obtain using just income data, as in (15) and (16)). With DWMD the two sets of estimates are similar because we are effectively putting more "identification weight" for the income shock variances on the income moments and less on the consumption moments (which display more sampling variability due to the imputation procedure).

of the permanent shock and the estimated variances of the transitory shock are generally higher in the second half of the 1980s (see also Figures 2 and 3). The MA parameter for the transitory shock is small. The variance of the imputation error σ_u^2 is always precisely measured and suggests that the imputation error absorbs a large amount of the cross-sectional variability in consumption in the PSID. The variance of heterogeneity in the consumption slope is also small but significant. In the whole sample the estimate of ϕ, the partial insurance coefficient for the permanent shock, provides evidence in favor of partial insurance. In particular, a 10 percent permanent income shocks induces a 6.1 percent permanent change in consumption. In contrast, the evidence on ψ accords with a simple PIH model with infinite horizon.[22] If we allow the partial insurance parameters to vary across time then we find a lower estimate for ϕ – indicating more insurance – in the later part of the 1980s. This is in line with the idea that a higher variance provides additional incentives to insure. However the differences in the partial insurance parameters over time are not statistically significant and hence we decided to restrict the coefficient to be constant over the whole period. The p-values for the test of constant insurance parameters over the two sub-periods are given in the last two rows of the table.[23]

There is much discussion in the literature on the reasons for the increase in income inequality of the last 25 years. In particular, there is much debate on whether the rise can be labeled permanent or transitory. In Figure 2 we plot the minimum distance estimate of the variance of the permanent shock, var($\zeta_{a,t}$), against time over the 1980s. There are two sets of estimates. One uses the full set of consumption and income moments for the baseline specification in Table VIII, and another just utilizes the income data. There is a close accordance between the two series which provides a check on the validity of our specification. The figure points to a strong growth in permanent income shocks during the early 1980s. The variance of permanent shocks levels off thereafter. This

[22]If we assume that food in the PSID reported in survey year t refers to that year rather than to the previous calendar year, we obtain similar results. The estimate of ϕ is slightly higher, but the qualitative pattern of results (and sensitivity checks) is unchanged.

[23]If we use a measure of consumption that includes the services from housing and vehicles we also obtain similar results (the estimate of ϕ is 0.53 with s.e. 0.10, and the estimate of ψ is 0.06 with s.e. 0.04).

evidence is similar to that reported by Moffitt and Gottschalk [1994] using PSID earnings data. It is also worth noting that from trough to peak the variance of the permanent shock doubles. A similar accordance between the two alternative estimates is also evident for the estimated variance of transitory shocks presented in Figure 3. This variance increases quite dramatically in the second half of the 1980s.

Table VIII also reports the results of the model for two education groups (with and without college education), and for two representative cohorts (born in the 1940s and in the 1930s). As before both the variance of the permanent shock and the variance of the transitory shock are generally higher in the second half of the 1980s.[24] The partial insurance parameter estimates point to interesting differences in insurance by type of household. In particular there appears to be more insurance in response to permanent shocks among the college educated group (indeed, we would not statistically reject the null hypothesis that there is *no* insurance in the group without college education). In contrast, the evidence on ψ accords with a simple PIH model and we cannot reject the null that there is full smoothing with respect to transitory shocks ($\psi = 0$) for both education groups. When the sample is stratified by year of birth, we find qualitatively similar results: there is evidence for full insurance with respect to transitory shocks and differences in the extent of insurance with respect to the permanent shocks. It is worth considering whether the presence of precautionary asset accumulation is stronger among older cohorts close to retirement. Recall that $\pi_{i,a,t}$ is the share of future labor income in the present value of lifetime wealth. Thus $\pi_{i,a,t}$ is likely to be lower for older cohort because older cohorts have both more accumulated financial wealth and lower prospective human capital wealth. We find evidence that permanent shocks are smoothed to a much greater extent than for younger cohorts. However, whether this is due to the effect played by precautionary wealth accumulation remarked above or by greater availability of insurance (such as social security or disability insurance) in the group of people born in the 1930s is something

[24] Since we stratify the sample by exogenous characteristics and estimate different parameters for different groups, we are effectively not considering the insurability of shocks across groups.

that we cannot address in the absence of additional information, such as panel data on assets and age-specific estimates of human capital wealth.

Having found evidence for partial insurance with respect to permanent income shocks, it is interesting trying to understand where it comes from. Table IX examines the impact of possible alternative insurance mechanisms. In particular, we focus on the insurance value of durable expenditures and of government taxes and transfers. Turning first to durables, one might expect the ϕ coefficient to rise simply because durables are more income elastic than nondurables. Moreover, with small costs of accessing the credit market (or small transaction costs in the second-hand market for durables), durable replacement could be used to smooth non-durable consumption in the face of income shocks, see Browning and Crossley [2003] for example. This would imply that with a measure of consumption that includes durables we should find less evidence for insurance, i.e., the estimated ϕ and ψ would rise. The second column of Table IX investigates this further. We use a comprehensive measure of consumption that includes durables and nondurables.[25] We repeat the imputation procedure detailed in Section 3.3, but use total consumption instead of nondurable consumption in the estimation of the demand equation (17). First we note that the ϕ coefficient does rise slightly as expected. Durables appear to provide some limited insurance for permanent shocks. Consider a permanent negative shock. In the absence of the durable hedge, one should reduce non durable consumption by the same amount of the shock. Downgrading one's car etc., and slowing the rate of replacement can help smoothing partially the non durable consumption effects of the permanent shock. A symmetric argument holds for a positive shock. However, note that durable expenditures appear to affect the size and statistical significance of the ψ coefficient to a much greater extent. This suggests that durables are particularly useful as a smoothing mechanism in response to transitory shocks. In this respect, they work as an imperfect form of savings as

[25] See Meyer and Sullivan [2001] for a detailed discussion of the measurement of durables in the CEX. Our measure of total consumption includes food, alcohol, tobacco, services, heating fuel, public and private transports (including gasoline), personal care, semidurables (clothing and footwear), and expenditure on durables, namely housing (mortgage interests, property tax, rents, other lodging, textiles, furniture, floor coverings, appliances), new and used cars, vehicle finance charges and insurance, car rentals and leases, health (insurance, prescription drugs, medical services), education, cash contributions, and personal insurance (life insurance and retirement).

suggested by Browning and Crossley (2003).

To see the impact of public insurance, suppose we exclude transfers (of any kind) from our measure of income. If taxes and transfers provide insurance for permanent income shocks, the insurance parameter in this specification should fall by an amount that reflects the degree of insurance. This happens because consumption still incorporates any insurance value of taxes and transfers but the new measure of income no longer does. The results of this experiment are reported in the last column of Table IX. A comparison with the baseline results shows that the estimated insurance parameter declines from 0.61 to 0.38. That is, by excluding transfers the partial insurance coefficient drops by a little over a third, an estimate of the insurance provided by private and public transfers. This insurance can also be seen through the change in the estimated variance of permanent and transitory shocks. With taxes and transfers excluded, the variances of income shocks are indeed much higher.[26] Note also that the estimate of ψ is barely affected, which seems to suggest that taxes and transfers help more in the smoothing of permanent shocks (disability insurance is one example) than in the smoothing of transitory shocks.

We next turn to our analysis of lower income families. Table X reports the results of extending our sample to the families of the SEO (the low-income subsample in the PSID). We present the estimates using the non-durable definition of consumption as in the baseline case and also the estimates including durable expenditures in the consumption definition. Two pieces of evidence are worth mentioning: the estimate of ϕ is higher reflecting less insurance opportunities in this sample, and we would now reject full insurance with respect to transitory shocks (an estimate of ψ of 0.12, not far from the 0.2 benchmark found by other researchers, as Hall and Mishkin [1982], who impute this excess sensitivity of consumption to transitory income shocks to binding liquidity constraints). Once we include durables we find a ϕ coefficient close to unity and the estimate of ψ rises to 0.224 indicating an appreciable degree of sensitivity, even to transitory shocks, among the

[26] This is a case where we reject the hypothesis that ϕ is the same in the two sub-periods 1979-84 and 1985-92. In practice, insurance tends to be higher in the second period perhaps because of the rise in the variance of earnings.

low income sample.

Finally, Figure 4 plots the variance of the permanent shock for various specifications and samples, including the whole sample (top left panel). Overall, trends in the variance are remarkably similar. One possible interpretation of this is that the differences in the estimates of ϕ that we find when we include the poverty sub-sample or durables, reflect genuine economic differences in access to insurance rather than differences in the variance of permanent shocks.

5 Conclusions

The aim of this paper has been to evaluate the degree of consumption insurance with respect to income shocks. This was achieved by investigating the degree to which the distribution of income shocks is transmitted to the distribution of consumption. For this we combined panel data on income from the PSID with consumption data from repeated CEX cross-sections. The framework allowed for self-insurance, in which consumers smooth idiosyncratic shocks through saving, and complete markets in which all idiosyncratic shocks are insured. Neither of these models were found to accord with the evidence.

We find some partial insurance for permanent shocks and almost complete insurance of transitory shocks. Only in low income households do we find significant sensitivity, and therefore only partial insurance, to transitory income shocks. Interestingly there appears to be a much greater degree of insurance of permanent shocks among the college educated. Not surprisingly there is also more insurance of such shocks for older cohorts. Our model suggests that we should see more insurance, even for permanent shocks, among those nearing retirement, especially where they have built up sufficient precautionary savings. The tax and welfare system are also found to play an important insurance role for permanent shocks.

When we include durables in our measure of consumption we find much less evidence of insurance of permanent shocks. The impact on the estimate of the insurance parameter for transitory income shocks is even greater. Here we find significant deviation from full insurance even for the regular

PSID sample. This effect gets stronger when we include the low income sub-sample. Durable expenditures, that is the timing and quality of durable replacement, appear to provide an insurance buffer between income shocks and non-durable expenditures. Especially for the lower income sample who may have less access or face higher transactions costs in the credit market.

Our results also show a strong growth in permanent income shocks in the US during the early 1980s (the variance of transitory shock also increases, but at a later stage). From trough to peak the variance of the permanent shock doubles, while the variance of the transitory shock only goes up by about 50%. The variance of permanent shocks levels off in the second half of the 1980s. The variance of the transitory shock is basically flat in the period where the variance of permanent shock is increasing, and it increases only when the variance of permanent shock slows down.

These results have implications for both macroeconomics and labor economics. The macroeconomic literature has long been concerned with explaining why modern economies depart from the complete markets benchmark. Recent work has examined the role of asymmetric information, moral hazard, heterogeneity, etc., and asked whether the complete markets model can be amended to include some form of imperfect insurance. This issue has not been subject to a systematic empirical investigation. Insofar as lack of smoothing opportunities implies a greater vulnerability to income shocks, our research can be relevant to issues of the incidence and permanence of poverty studied in the labor economics literature. Studying how well families smooth income shocks, how this changes over time in response to changes in the economic environment confronted, and how different household types differ in their smoothing opportunities, is an important complement to understanding the effect of redistributive policies and anti-poverty strategies.

References

[1] Abowd, J., and D. Card (1989), "On the covariance structure of earnings and hours changes", *Econometrica*, **57**, 411-45.

[2] Aiyagari, R. (1994), "Uninsured risk and aggregate saving", *Quarterly Journal of Economics*, **109**, 659-84

[3] Altonji, J., A.P. Martins, and A. Siow (2002), "Dynamic factor models of consumption, hours, and income", *Research in Economics,* **56**, 3-59.

[4] Altonji, J., and L. Segal (1996), "Small-sample bias in GMM estimation of covariance structures", *Journal of Business and Economic Statistics,* **14**, 353-66.

[5] Altonji, J., and A. Siow (1987), "Testing the response of consumption to income changes with (noisy) panel data", *Quarterly Journal of Economics*, **102**, 293-28.

[6] Alvarez, F. and U. Jermann (2000), "Efficiency, equilibrium and asset pricing with risk of default", *Econometrica*, **68**, 775-97.

[7] Attanasio, O., E. Battistin and H. Ichimura (2004), "What really happened to consumption inequality in the US?", University College London, *mimeo.*

[8] Attanasio, O., and S. Davis (1996), "Relative wage movements and the distribution of consumption", *Journal of Political Economy,* **104**, 1227-62.

[9] Attanasio, O., and V. Rios Rull (2000), "Consumption smoothing in island economies: Can public insurance reduce welfare?", *European Economic Review*, **44**, 1225-58.

[10] Attanasio, O., and G. Weber (1995), "Is consumption growth consistent with intertemporal optimization: Evidence from the Consumer Expenditure Survey", *Journal of Political Economy*, **103**, 1121-57.

[11] Auerbach, A.J., and D. Feenberg (2000), "The significance of Federal Taxes as automatic stabilizers", *Journal of Economic Perspectives*, **14** (Summer), 37-56.

[12] Blundell, R., H. Low and I. Preston (2004), "Income Risk and Consumption Inequality: A Simulation Study", Institute for Fiscal Studies WP04/26 (http://www.ifs.org.uk/workingpapers/wp0426.pdf).

[13] Blundell, R., and L. Pistaferri (2003), "Income volatility and household consumption: The impact of food assistance programs", *Journal of Human Resources,* Vol 38, 1032-1050.

[14] Blundell, R., L. Pistaferri and I. Preston (2004), "Imputing consumption in the PSID using food demand estimates from the CEX", Institute for Fiscal Studies WP04/27 (http://www.ifs.org.uk/workingpapers/wp0427.pdf).

[15] Blundell, R., and I. Preston (1998), "Consumption inequality and income uncertainty", *Quarterly Journal of Economics* **113**, 603-640.

[16] Blundell, R., and T. Stoker (2004), "Aggregation and Heterogeneity", forthcoming in *Journal of Economic Literature.*

[17] Browning, M., and T. Crossley (2003), "Shocks, stocks and socks: Consumption smoothing and the replacement of durables", McMaster Economics Working Paper 2003-7.

[18] Carroll, C. (2001), "Precautionary saving and the marginal propensity to consume out of permanent income", NBER Working Paper 8233.

[19] Chamberlain, G. (1984), "Panel data", in *Handbook of Econometrics*, vol. 2, edited by Zvi Griliches and Michael D. Intriligator. Amsterdam: North-Holland.

[20] Cutler, D., and L. Katz (1992), "Rising inequality? Changes in the distribution of income and consumption in the 1980s", *American Economic Review,* **82**, 546-51.

[21] Deaton, A. (1992), *Understanding Consumption.* Baltimore: John Hopkins University Press.

[22] Deaton, A. (1997), *The Analysis of Household Surveys.* Baltimore: John Hopkins University Press.

[23] Deaton, A., and C. Paxson (1994), "Intertemporal choice and inequality", *Journal of Political Economy,* **102**, 384-94.

[24] Dynarski, S., and J. Gruber (1997), "Can families smooth variable earnings?", *Brooking Papers on Economic Activity*, **1**, 229-305.

[25] Engen, E., and J. Gruber (2001), "Unemployment insurance and precautionary savings", *Journal of Monetary Economics*, 47, 545-79.

[26] Fay, S., E. Hurst, and M. White (2002), "The consumer bankruptcy decision", *American Economic Review*, **92**, 706-18.

[27] Flavin, M. A. (1981), "The Adjustment of Consumption to Changing Expectations about Future Income," *Journal of Political Economy*, vol. 89, no.5.

[28] Gruber, J., and A. Yelowitz (1999), "Public Health Insurance and Private Savings," *Journal of Political Economy*, 107(6), 1249-74.

[29] Hall, R., and F. Mishkin (1982), "The sensitivity of consumption to transitory income: Estimates from panel data of households", *Econometrica,* **50**, 261-81.

[30] Hayashi, F., J. Altonji, and L. Kotlikoff (1996), "Risk sharing between and within families", *Econometrica*, **64**, 261-94.

[31] Heathcote, J., K. Storesletten and G. Violante (2004), "The Macroeconomic Implications of Rising Wage Inequality in the United States", Working Paper.

[32] Hill, M. (1992), *The Panel Study of Income Dynamics: A user's guide*, Newbury Park, California: Sage Publications.

[33] Hurst, E., and F. Stafford (2003), "Home is where the equity is: Liquidity constraints, refinancing and consumption", *Journal of Money, Credit, and Banking*, forthcoming.

[34] Jappelli, T., and L. Pistaferri (2004), "Intertemporal choice and consumption mobility", University of Salerno and Stanford University, *mimeo.*

[35] Kehoe, T., and D. Levine (2001), "Liquidity constrained markets versus debt constrained markets", *Econometrica*, **69**, 575-98

[36] Kniesner, T., and J. Ziliak (2002), "Tax reform and automatic stabilization", *American Economic Review*, **92**, 590-612.

[37] Kotlikoff, L., and A. Spivak (1981), "The family as an incomplete annuities market", *Journal of Political Economy*, **89**, 372-91.

[38] Krueger, D., and F. Perri (2003), "Does Income Inequality Lead to Consumption Inequality? Evidence and Theory", Working paper.

[39] MaCurdy, T. (1982), "The use of time series processes to model the error structure of earnings in a longitudinal data analysis", *Journal of Econometrics*, **18**, 82-114.

[40] Mankiw, M., and M. Kimball (1992), "Precautionary saving and the timing of taxes", *Journal of Political Economy*, **97**, 863-79.

[41] Meghir, C., and L. Pistaferri (2004), "Income variance dynamics and heterogeneity", *Econometrica*, 72(1), 1-32.

[42] Meyer, B. and J. Sullivan (2001), "The living conditions of single mothers in the 19080s and 1990s", NBER Working Paper 8298, May.

[43] Moffitt, R., and P. Gottschalk (1994), "Trends in the autocovariance structure of earnings in the US: 1969-1987", Brown University, *mimeo.*

[44] Moffitt, R., and P. Gottschalk (2002), "Trends in the Transitory Variance of Earnings in the United States", *Economic Journal*, vol. 112(127), pages C68-C73, March.

[45] Skinner, J. (1987), "A superior measure of consumption from the Panel Study of Income Dynamics", *Economic Letters*, **23**, 213-16.

[46] Stephens, M. Jr. (2002), "Worker displacement and the added worker effect", *Journal of Labor Economics,* **20**, 504-37.

Table I
Sample selection in the PSID

	# dropped	# remain
Initial sample (1968-1992)	0	145,940
Interviewed prior to 1978	52,408	93,532
Change in family composition	18,570	74,962
Female head	23,779	51,183
Missing values and topcoding	308	50,875
Change in marital status	5,882	44,993
Income outliers	2,407	42,586
Born before 1920 or after 1959	8,510	34,076
Poverty subsample	12,600	21,476
Aged less than 30 or more than 65	3,674	17,778

Table II
Sample selection in the CEX

	# dropped	# remain
Initial sample	0	141,289
Missing expenditure data	1,351	139,938
Present for less than 12 months	76,773	63,165
Observed after 1992	19,310	43,855
Zero before-tax income	1,308	42,547
Missing region or education	14,029	28,418
Marital status	5,848	22,570
Born before 1920 or after 1959	4,648	17,922
Aged less than 30 or more than 65	1,843	16,079
Income outliers and incomplete income response	942	15,137

Table III
Comparison of means, PSID and CEX

	1980		1983		1986		1989		1992	
	PSID	CEX	PSID	CEX	PSID	CEX	PSID	CEX	PSID	CEX
Age	42.97	43.58	43.36	44.90	43.84	46.01	44.00	45.26	45.89	47.01
Family size	3.61	3.98	3.52	3.74	3.48	3.64	3.44	3.60	3.42	3.55
# of children	1.31	1.49	1.25	1.28	1.21	1.19	1.18	1.17	1.14	1.15
White	0.91	0.89	0.92	0.88	0.92	0.89	0.93	0.89	0.93	0.88
HS dropout	0.21	0.20	0.17	0.19	0.16	0.18	0.14	0.14	0.13	0.15
HS graduate	0.30	0.33	0.31	0.33	0.32	0.30	0.32	0.30	0.31	0.30
College dropout	0.49	0.47	0.52	0.48	0.53	0.53	0.54	0.56	0.56	0.55
Northeast	0.21	0.20	0.21	0.24	0.22	0.21	0.22	0.23	0.22	0.22
Midwest	0.33	0.28	0.31	0.26	0.30	0.27	0.30	0.28	0.30	0.29
South	0.31	0.28	0.31	0.28	0.30	0.27	0.31	0.27	0.30	0.26
West	0.15	0.24	0.17	0.21	0.18	0.25	0.18	0.23	0.18	0.23
Husband working	0.96	0.97	0.94	0.92	0.93	0.90	0.94	0.92	0.93	0.88
Wife working	0.69	0.67	0.71	0.66	0.74	0.71	0.78	0.72	0.77	0.73
Family income	32,759	29,078	37,907	35,923	45,035	43,630	52,919	51,205	61,911	56,520
Food expenditure	4,449	4,656	4,858	4,617	5,306	5,199	5,864	6,135	6,620	6,431

Table IV
The demand for food in the CEX

This table reports IV estimates of the demand equation for (the logarithm of) food spending in the CEX. We instrument the log of total nondurable expenditure (and its interaction with time, education, and kids dummies) with the cohort-education-year specific average of the log of the husband's hourly wage and the cohort-education-year specific average of the log of the wife's hourly wage (and their interactions with time, education and kids dummies). Standard errors are in round parenthesis; the Shea's partial R^2 for the relevance of instruments in square brackets. In all cases, the p-value of the F-test on the excluded instrument is <0.01 percent.

Variable	Estimate	Variable	Estimate	Variable	Estimate
$\ln c$	0.8791 (0.1613) [0.011]	$\ln c * 1992$	0.0040 (0.0054) [0.090]	Family size	0.0244 (0.0097)
$\ln c$*High School dropout	0.0356 (0.0714) [0.050]	$\ln c$*One child	0.0215 (0.0344) [0.142]	$\ln p_{food}$	−0.9642 (0.2220)
$\ln c$*High School graduate	0.0821 (0.0913) [0.026]	$\ln c$*Two children	−0.0205 (0.0370) [0.128]	$\ln p_{transports}$	7.1099 (7.8560)
$\ln c * 1981$	0.1152 (0.1064) [0.056]	$\ln c$*Three children+	0.0318 (0.0345) [0.188]	$\ln p_{fuel+utils}$	−1.7335 (4.0720)
$\ln c * 1982$	0.0630 (0.0813) [0.054]	One child	−0.1644 (0.3330)	$\ln p_{alcohol+tobacco}$	−2.2015 (4.5968)
$\ln c * 1983$	0.0562 (0.0697) [0.049]	Two children	0.2956 (0.3569)	Born 1955-59	−0.0768 (0.0561)
$\ln c * 1984$	0.0564 (0.0660) [0.051]	Three children+	−0.1889 (0.3350)	Born 1950-54	−0.0333 (0.0484)
$\ln c * 1985$	0.0378 (0.0640) [0.063]	High school dropout	−0.3596 (0.6751)	Born 1945-49	−0.0316 (0.0414)
$\ln c * 1986$	0.0287 (0.0589) [0.067]	High school graduate	−0.8819 (0.8574)	Born 1940-44	−0.0241 (0.0352)
$\ln c * 1987$	0.0637 (0.0589) [0.066]	Age	0.0081 (0.0091)	Born 1935-39	−0.0229 (0.0278)
$\ln c * 1988$	0.0469 (0.0450) [0.050]	Age^2	−0.0001 (0.0001)	Born 1930-34	−0.0134 (0.0197)
$\ln c * 1989$	0.0409 (0.0366) [0.047]	Northeast	0.0008 (0.0073)	Born 1925-29	−0.0133 (0.0144)
$\ln c * 1990$	0.0195 (0.0293) [0.060]	Midwest	−0.0253 (0.0095)	White	0.0795 (0.0130)
$\ln c * 1991$	−0.0029 (0.0313) [0.112]	South	−0.0308 (0.0091)	Constant	−0.9157 (0.9881)
OID test				16.407 (d.f. 18; p-value 56%)	

Table V
The autocovariance matrix of income growth

Year	var(Δy_t)	cov($\Delta y_{t+1}, \Delta y_t$)	cov($\Delta y_{t+2}, \Delta y_t$)
1980	0.0830	-0.0224	-0.0019
	(0.0088)	(0.0041)	(0.0030)
1981	0.0813	-0.0291	-0.0038
	(0.0090)	(0.0049)	(0.0035)
1982	0.0784	-0.0231	-0.0059
	(0.0064)	(0.0039)	(0.0029)
1983	0.0859	-0.0242	-0.0093
	(0.0092)	(0.0041)	(0.0053)
1984	0.0861	-0.0310	-0.0028
	(0.0059)	(0.0038)	(0.0038)
1985	0.0927	-0.0321	-0.0012
	(0.0069)	(0.0053)	(0.0042)
1986	0.1153	-0.0440	-0.0078
	(0.0120)	(0.0094)	(0.0061)
1987	0.1185	-0.0402	0.0014
	(0.0115)	(0.0052)	(0.0046)
1988	0.0929	-0.0313	-0.0017
	(0.0084)	(0.0041)	(0.0032)
1989	0.0921	-0.0303	-0.0026
	(0.0071)	(0.0075)	(0.0042)
1990	0.0988	-0.0303	-0.0056
	(0.0135)	(0.0056)	(0.0045)
1991	0.1136	-0.0409	NA
	(0.0245)	(0.0141)	
1992	0.1279	NA	NA
	(0.0153)		

Table VI
The autocovariance matrix of consumption growth

Year	var(Δc_t)	cov($\Delta c_{t+1}, \Delta c_t$)	cov($\Delta c_{t+2}, \Delta c_t$)
1980	0.1266	-0.0573	0.0020
	(0.0106)	(0.0087)	(0.0053)
1981	0.1172	-0.0545	0.0029
	(0.0115)	(0.0072)	(0.0043)
1982	0.1241	-0.0587	0.0005
	(0.0099)	(0.0080)	(0.0049)
1983	0.1386	-0.0634	-0.0018
	(0.0130)	(0.0069)	(0.0059)
1984	0.1559	-0.0738	-0.0121
	(0.0127)	(0.0117)	(0.0082)
1985	0.1719	-0.0823	NA
	(0.0214)	(0.0188)	
1986	0.1528	NA	NA
	(0.0177)		
1987	NA	NA	NA
1988	NA	NA	NA
1989	NA	NA	NA
1990	0.1604	-0.0579	-0.0061
	(0.0196)	(0.0057)	(0.0064)
1991	0.1457	-0.0619	NA
	(0.0099)	(0.0083)	
1992	0.1371	NA	NA
	(0.0115)		

Table VII
The consumption-income growth covariance matrix

Year	cov($\Delta y_t, \Delta c_t$)	cov($\Delta y_t, \Delta c_{t+1}$)	cov($\Delta y_{t+1}, \Delta c_t$)
1980	0.0039	0.0052	0.0012
	(0.0040)	(0.0036)	(0.0037)
1981	0.0102	-0.0052	-0.0050
	(0.0036)	(0.0035)	(0.0033)
1982	0.0162	-0.0016	-0.0056
	(0.0037)	(0.0040)	(0.0032)
1983	0.0204	-0.0055	-0.0075
	(0.0044)	(0.0042)	(0.0046)
1984	0.0221	-0.0105	-0.0054
	(0.0048)	(0.0044)	(0.0044)
1985	0.0177	-0.0033	-0.0024
	(0.0063)	(0.0063)	(0.0054)
1986	0.0165	NA	0.0001
	(0.0048)		(0.0051)
1987	NA	NA	NA
1988	NA	NA	NA
1989	NA	0.0039	NA
		(0.0039)	
1990	0.0067	-0.0033	0.0026
	(0.0042)	(0.0039)	(0.0060)
1991	0.0101	-0.0073	-0.0000
	(0.0040)	(0.0040)	(0.0042)
1992	0.0111	NA	NA
	(0.0038)		

Table VIII
Minimum distance partial insurance and variance estimates

This table reports DWMD results of the parameters of interest. See the main text for details. Standard errors in parenthesis.

		Whole sample	No College	College	Born 1940s	Born 1930s
σ^2_ζ	1979-81	0.0078 (0.0037)	0.0054 (0.0045)	0.0077 (0.0058)	0.0066 (0.0040)	0.0074 (0.0072)
	1982	0.0204 (0.0051)	0.0154 (0.0065)	0.0275 (0.0085)	0.0207 (0.0063)	0.0195 (0.0120)
	1983	0.0316 (0.0068)	0.0348 (0.0110)	0.0269 (0.0096)	0.0212 (0.0062)	0.0208 (0.0092)
	1984	0.0298 (0.0051)	0.0340 (0.0085)	0.0244 (0.0070)	0.0214 (0.0077)	0.0205 (0.0104)
	1985	0.0261 (0.0063)	0.0372 (0.0090)	0.0133 (0.0085)	0.0206 (0.0074)	0.0206 (0.0123)
	1986	0.0250 (0.0076)	0.0247 (0.0094)	0.0324 (0.0127)	0.0218 (0.0115)	0.0075 (0.0167)
	1987	0.0248 (0.0070)	0.0178 (0.0080)	0.0308 (0.0112)	0.0187 (0.0110)	0.0128 (0.0163)
	1988	0.0141 (0.0069)	0.0215 (0.0121)	0.0069 (0.0069)	0.0100 (0.0074)	0.0300 (0.0245)
	1989	0.0218 (0.0069)	0.0166 (0.0117)	0.0316 (0.0069)	0.0175 (0.0089)	0.0337 (0.0196)
	1990-92	0.0145 (0.0047)	0.0096 (0.0049)	0.0283 (0.0081)	0.0115 (0.0074)	0.0099 (0.0108)
σ^2_ε	1979	0.0415 (0.0058)	0.0510 (0.0095)	0.0325 (0.0057)	0.0292 (0.0051)	0.0376 (0.0083)
	1980	0.0319 (0.0043)	0.0332 (0.0058)	0.0311 (0.0064)	0.0281 (0.0059)	0.0281 (0.0066)
	1981	0.0373 (0.0054)	0.0378 (0.0057)	0.0365 (0.0093)	0.0328 (0.0060)	0.0450 (0.0156)
	1982	0.0287 (0.0044)	0.0370 (0.0072)	0.0211 (0.0046)	0.0259 (0.0050)	0.0276 (0.0064)
	1983	0.0288 (0.0044)	0.0393 (0.0080)	0.0195 (0.0041)	0.0200 (0.0051)	0.0301 (0.0074)
	1984	0.0352 (0.0042)	0.0402 (0.0063)	0.0319 (0.0056)	0.0219 (0.0048)	0.0512 (0.0117)
	1985	0.0381 (0.0051)	0.0320 (0.0061)	0.0445 (0.0082)	0.0278 (0.0067)	0.0507 (0.0115)
	1986	0.0546 (0.0103)	0.0445 (0.0080)	0.0625 (0.0176)	0.0543 (0.0247)	0.0742 (0.0186)
	1987	0.0481 (0.0054)	0.0531 (0.0083)	0.0451 (0.0073)	0.0496 (0.0102)	0.0544 (0.0112)
	1988	0.0384 (0.0045)	0.0425 (0.0066)	0.0354 (0.0063)	0.0398 (0.0069)	0.0509 (0.0131)
	1989	0.0380 (0.0081)	0.0501 (0.0144)	0.0224 (0.0061)	0.0369 (0.0080)	0.0407 (0.0111)
	1990-92	0.0474 (0.0057)	0.0488 (0.0055)	0.0345 (0.0063)	0.0402 (0.0065)	0.0691 (0.0117)
θ		0.0884 (0.0273)	0.1011 (0.0346)	0.0845 (0.0363)	0.0999 (0.0434)	0.1137 (0.0518)
$\sigma^2_{u^c}$		0.0600 (0.0030)	0.0725 (0.0052)	0.0469 (0.0029)	0.0547 (0.0046)	0.0577 (0.0057)
σ^2_ξ		0.0114 (0.0036)	0.0112 (0.0063)	0.0139 (0.0039)	0.0146 (0.0060)	0.0166 (0.0064)
ϕ		0.6100 (0.1107)	0.7980 (0.2176)	0.3655 (0.0894)	0.7266 (0.2096)	0.5394 (0.2476)
ψ		0.0519 (0.0368)	0.0859 (0.0514)	0.0361 (0.0503)	0.0851 (0.0642)	0.0278 (0.0577)
P-value test of equal ϕ		91%	66%	4%	15%	56%
P-value test of equal ψ		37%	42%	20%	45%	13%

Table IX
Minimum distance partial insurance and variance estimates

This table reports DWMD results of the parameters of interest. See the main text for details. Standard errors in parenthesis.

Consumption:		Non dur.	Total	Non dur.
Income:		Earn.+transf.	Earn.+transf.	Earnings only
Sample:		Baseline	Baseline	Baseline
σ^2_ζ	1979-81	0.0078 (0.0037)	0.0176 (0.0052)	0.0102 (0.0060)
	1982	0.0204 (0.0051)	0.0225 (0.0054)	0.0390 (0.0086)
	1983	0.0316 (0.0068)	0.0353 (0.0073)	0.0656 (0.0146)
	1984	0.0298 (0.0051)	0.0295 (0.0054)	0.0641 (0.0146)
	1985	0.0261 (0.0063)	0.0267 (0.0068)	0.0542 (0.0209)
	1986	0.0250 (0.0076)	0.0282 (0.0094)	0.0522 (0.0186)
	1987	0.0248 (0.0070)	0.0264 (0.0070)	0.0475 (0.0159)
	1988	0.0141 (0.0069)	0.0149 (0.0070)	0.0443 (0.0187)
	1989	0.0218 (0.0069)	0.0250 (0.0072)	0.0426 (0.0147)
	1990-92	0.0145 (0.0047)	0.0249 (0.0070)	0.0434 (0.0116)
σ^2_ε	1979	0.0415 (0.0058)	0.0366 (0.0061)	0.0523 (0.0078)
	1980	0.0319 (0.0043)	0.0289 (0.0044)	0.0435 (0.0068)
	1981	0.0373 (0.0054)	0.0346 (0.0054)	0.0471 (0.0071)
	1982	0.0287 (0.0044)	0.0278 (0.0044)	0.0426 (0.0077)
	1983	0.0288 (0.0044)	0.0281 (0.0044)	0.0716 (0.0169)
	1984	0.0352 (0.0042)	0.0352 (0.0042)	0.0604 (0.0106)
	1985	0.0381 (0.0051)	0.0373 (0.0053)	0.0646 (0.0099)
	1986	0.0546 (0.0103)	0.0528 (0.0105)	0.1211 (0.0332)
	1987	0.0481 (0.0054)	0.0473 (0.0054)	0.0648 (0.0107)
	1988	0.0384 (0.0045)	0.0377 (0.0045)	0.0840 (0.0143)
	1989	0.0380 (0.0081)	0.0349 (0.0083)	0.0700 (0.0140)
	1990-92	0.0474 (0.0057)	0.0435 (0.0059)	0.0820 (0.0112)
θ		0.0884 (0.0273)	0.0810 (0.0295)	0.0688 (0.0309)
$\sigma^2_{u^c}$		0.0600 (0.0030)	0.1879 (0.0126)	0.0600 (0.0030)
σ^2_ξ		0.0114 (0.0036)	0.0399 (0.0136)	0.0123 (0.0035)
ϕ		0.6100 (0.1107)	0.7324 (0.1318)	0.3780 (0.0725)
ψ		0.0519 (0.0368)	0.1720 (0.0843)	0.0552 (0.0298)
P-value test of equal ϕ		91%	20%	0.4%
P-value test of equal ψ		37%	73%	13%

Table X

Minimum distance partial insurance and variance estimates

This table reports DWMD results of the parameters of interest. See the main text for details. Standard errors in parenthesis.

Consumption:		Non dur.	Non dur.	Total
Income:		Earn.+transf.	Earn.+transf.	Earn.+transf.
Sample:		Baseline	Baseline+SEO	Baseline+SEO
σ^2_ζ	1979-81	0.0078 (0.0037)	0.0119 (0.0036)	0.0176 (0.0041)
	1982	0.0204 (0.0051)	0.0235 (0.0047)	0.0219 (0.0047)
	1983	0.0316 (0.0068)	0.0301 (0.0056)	0.0260 (0.0056)
	1984	0.0298 (0.0051)	0.0319 (0.0049)	0.0287 (0.0052)
	1985	0.0261 (0.0063)	0.0229 (0.0049)	0.0213 (0.0051)
	1986	0.0250 (0.0076)	0.0216 (0.0056)	0.0208 (0.0059)
	1987	0.0248 (0.0070)	0.0199 (0.0050)	0.0201 (0.0050)
	1988	0.0141 (0.0069)	0.0168 (0.0052)	0.0171 (0.0053)
	1989	0.0218 (0.0069)	0.0199 (0.0053)	0.0222 (0.0054)
	1990-92	0.0145 (0.0047)	0.0190 (0.0043)	0.0292 (0.0053)
σ^2_ε	1979	0.0415 (0.0058)	0.0468 (0.0047)	0.0445 (0.0048)
	1980	0.0319 (0.0043)	0.0373 (0.0046)	0.0356 (0.0046)
	1981	0.0373 (0.0054)	0.0416 (0.0044)	0.0402 (0.0044)
	1982	0.0287 (0.0044)	0.0316 (0.0040)	0.0327 (0.0040)
	1983	0.0288 (0.0044)	0.0428 (0.0063)	0.0443 (0.0062)
	1984	0.0352 (0.0042)	0.0418 (0.0039)	0.0424 (0.0039)
	1985	0.0381 (0.0051)	0.0443 (0.0042)	0.0446 (0.0042)
	1986	0.0546 (0.0103)	0.0534 (0.0073)	0.0533 (0.0073)
	1987	0.0481 (0.0054)	0.0514 (0.0050)	0.0512 (0.0050)
	1988	0.0384 (0.0045)	0.0473 (0.0046)	0.0470 (0.0046)
	1989	0.0380 (0.0081)	0.0432 (0.0063)	0.0409 (0.0065)
	1990-92	0.0474 (0.0057)	0.0503 (0.0043)	0.0465 (0.0044)
θ		0.0884 (0.0273)	0.0832 (0.0211)	0.0807 (0.0218)
$\sigma^2_{u^c}$		0.0600 (0.0030)	0.0763 (0.0034)	0.2025 (0.0108)
σ^2_ξ		0.0114 (0.0036)	0.0198 (0.0044)	0.0833 (0.0148)
ϕ		0.6100 (0.1107)	0.6974 (0.1037)	1.0629 (0.1584)
ψ		0.0519 (0.0368)	0.1235 (0.0328)	0.2239 (0.0649)
P-value test of equal ϕ		91%	58%	0.2%
P-value test of equal ψ		37%	87%	23%

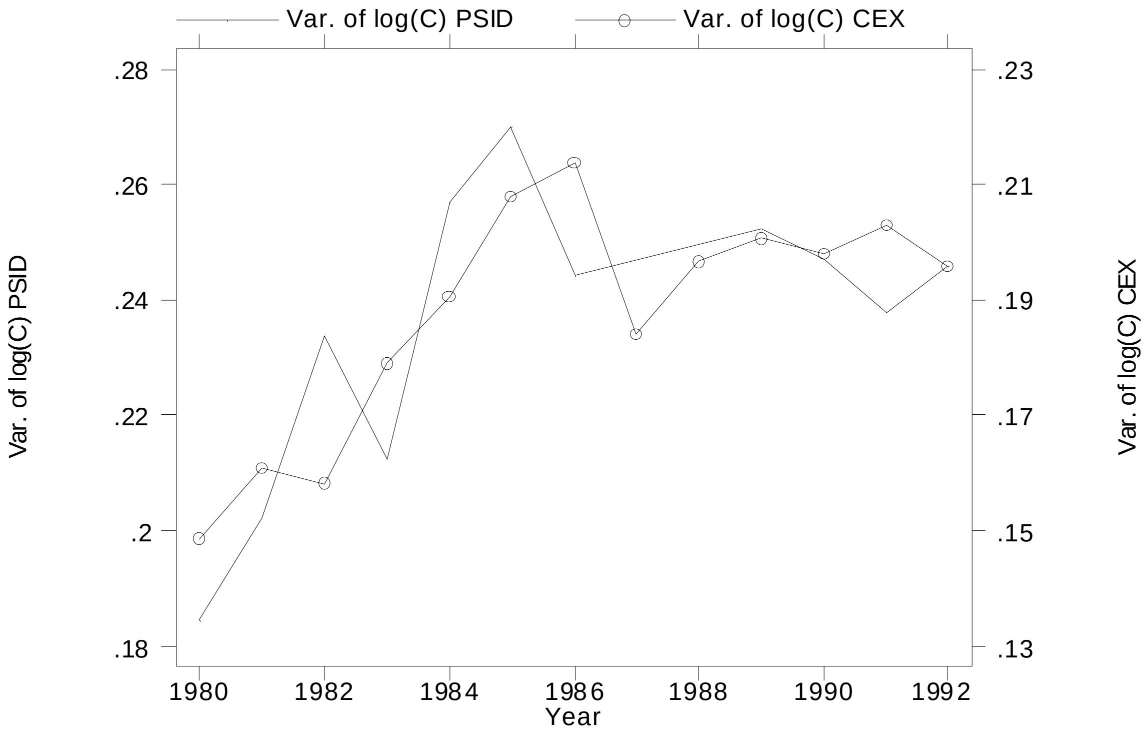

Figure 1: The Variance of Consumption, PSID and CEX.

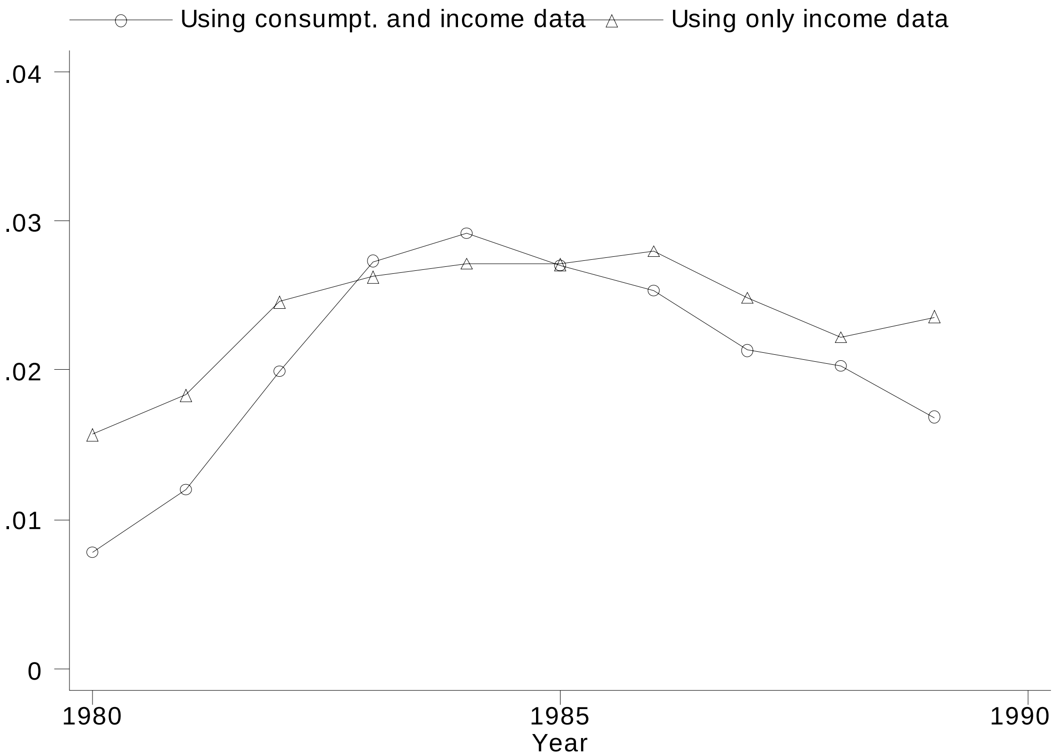

Figure 2: The Variance of the Permanent Shock in the 1980s.

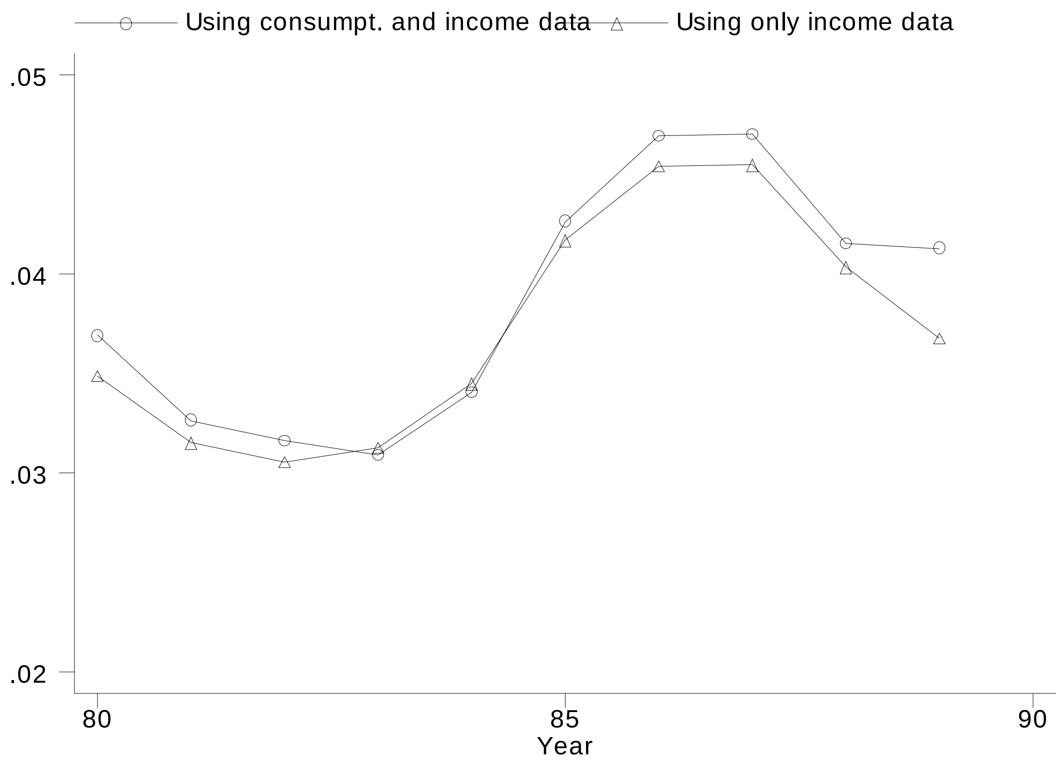

Figure 3: The Variance of the Transitory Shock in the 1980s

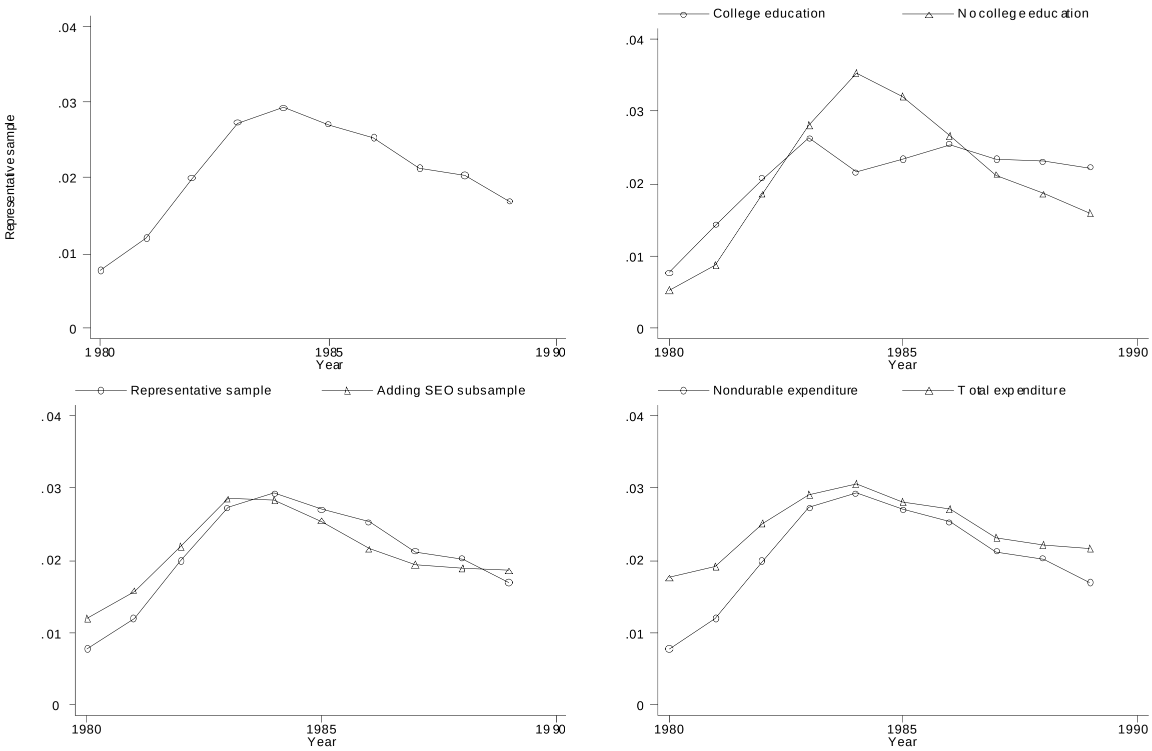

Figure 4: The variance of permanent shocks in various specifications and samples.

A.1 Appendix: The Euler Equation Approximation

If preferences are quadratic (and interest rates are not subject to uncertainty), it is possible to obtain a closed form solution for consumption. It is also straightforward to derive an exact mapping between the expectation error of the Euler equation for consumption and income shocks. See Hall and Mishkin [1982] for example. Quadratic preferences have well known undesirable features, such as increasing risk aversion and lack of a precautionary motive for saving. More realistic preferences, such as the CRRA functional form used here, solve these problems but deliver no closed form solution for consumption. The Euler equation can be linearized to describe the behavior of consumption growth. Moreover, in this appendix we derive an approximation of the mapping between the expectation error of the Euler equation and the income shock.

Consider approximating the logarithm of the sum of an arbitrary series $X_t,\ X_{t+1},\dots,X_S$:

$$\ln\sum_{k=0}^{S-t}X_{t+k}=\ln X_t+\ln\left[1+\sum_{k=1}^{S-t}\exp(\ln X_{t+k}-\ln X_t)\right]$$

Taking a Taylor expansion around $\ln X_{t+k}=\ln X_t+\sum_{i=0}^{k}\delta_{t+i}$, $k=1,...,S-t$ for some path of increments $\delta_t,\ \delta_{t+1},\dots,\delta_S$ with $\delta_t=0$,

$$\begin{aligned}\ln\sum_{k=0}^{S-t}X_{t+k}&\simeq\ln X_t+\ln\left[1+\sum_{k=1}^{S-t}\exp(\sum_{i=0}^{k}\delta_{t+i})\right]\\&\quad+\sum_{k=1}^{S-t}\frac{\exp(\sum_{i=0}^{k}\delta_{t+i})}{\left[1+\sum_{k=1}^{S-t}\exp(\sum_{i=0}^{k}\delta_{t+i})\right]}(\ln X_{t+k}-\ln X_t-\sum_{i=0}^{k}\delta_{t+i})\\&\simeq\sum_{k=0}^{S-t}\alpha_{t+k,S}^{\delta}[\ln X_{t+k}-\ln\alpha_{t+k,S}^{\delta}]\end{aligned}\tag{18}$$

where $\alpha_{t+k,S}^{\delta}=\exp(\sum_{i=0}^{k}\delta_{t+i})/\left[1+\sum_{k=1}^{S-t}\exp(\sum_{i=0}^{k}\delta_{t+i})\right]$.

Take the consumption income model as in (4). With CRRA preferences, optimization implies the Euler equation

$$C_{i,a-1,t-1}^{\beta-1}=(1+r_{t-1})e^{\Delta Z'_{i,a,t}\vartheta_t}E_{a-1,t-1}C_{i,a,t}^{\beta-1}$$

and therefore, approximately,

$$\Delta\log C_{i,a,t}\simeq\Delta Z'_{i,a,t}\vartheta_t+\eta_{i,a,t}+\Omega_{i,a,t}$$

where $\eta_{i,a,t}$is a consumption shock with $E_{a-1,t-1}\eta_{i,a,t}=0$ and $\Omega_{i,a,t}$ captures any slope in the consumption path due to interest rates, impatience or precautionary savings. Suppose that any idiosyncratic component to this gradient to the consumption path can be adequately picked up by a cohort/time-specific component $\Gamma_{b,t}$ and an individual element $\xi_{i,a,t}$

$$\Delta\log C_{i,a,t}-\Gamma_{b,t}-\Delta Z'_{i,a,t}\vartheta_t=\Delta c_{i,a,t}\simeq\eta_{i,a,t}+\xi_{i,a,t}.$$

From (3) we also have

$$\Delta y_{i,a+k,t+k}=\zeta_{i,a+k,t+k}+\sum_{j=0}^{q}\theta_j\varepsilon_{i,a+k-j,t+k-j}.$$

The intertemporal budget constraint is

$$\sum_{k=0}^{T-t}Q_{t+k}C_{t+k}=\sum_{k=0}^{L-t}Q_{t+k}Y_{t+k}+A_t$$

where T is death, L is retirement and Q_{t+k} is appropriate discount factor $\prod_{i=1}^{k}(1+r_{t+i})$, $k=1,...,T-t$ (and $Q_t=1$).

Applying the approximation (18) appropriately to each side

$$\begin{aligned}
&\sum_{k=0}^{T-t} \alpha_{t+k,T}^{\Omega+\Delta Z\vartheta-r}[\ln C_{t+k} - \ln Q_{t+k} - \ln \alpha_{t+k,T}^{\omega-r}] \\
&\simeq \pi_{i,a,t} \sum_{k=0}^{L-t} \alpha_{t+k,L}^{\Delta Z\varphi-r}[\ln Y_{t+k} - \ln Q_{t+k} - \ln \alpha_{t+k,L}^{\Delta Z\varphi-r}] \\
&+ (1-\pi_{i,a,t}) \ln A_t - [(1-\pi_{i,a,t}) \ln(1-\pi_{i,a,t}) + \pi_{i,a,t} \ln \pi_{i,a,t}]
\end{aligned}$$

where $\pi_{i,a,t} = \frac{\sum_{k=0}^{L-t} Q_{t+k}Y_{t-k}}{\sum_{k=0}^{L-t} Q_{t+k}Y_{t-k}+A_t}$ is the share of future labor income in current human and financial wealth.

Taking differences in expectations gives

$$\eta_{i,a,t} \simeq \pi_{i,a,t}\left[\zeta_{i,a,t} + \gamma_{t,L}\varepsilon_{i,a,t}\right]$$

where $\gamma_{t,L} = \left(\sum_{j=0}^{q} \alpha_{t+j,L}^{\Delta Z\vartheta-r}\theta_j\right)$.

Then[27]

$$\Delta c_{i,a,t} \simeq \xi_{i,a,t} + \pi_{i,a,t}\zeta_{i,a,t} + \gamma_{t,L}\pi_{i,a,t}\varepsilon_{i,a,t}.$$

If $\Delta Z'_{i,t+k}\vartheta_t = 0$ and $r_t = r$ is constant then $\alpha_{t+j,L}^{\Delta Z\vartheta-r} = \alpha_{t+j,L}^{-r} = \exp(-jr)/\sum_{k=0}^{L}\exp(-kr) \simeq r/(1+r)^k$ and $\gamma_{t,L} \simeq \frac{r}{1+r}[1+\sum_{j=1}^{q}\theta_j/(1+r)^j]$.

A.2 Appendix: Measurement error

In the light of our imputation procedure, let's assume that both consumption and income are measured with multiplicative independent error, e.g., $y_{i,a,t}^* = y_{i,a,t} + u_{i,a,t}^y$ and $c_{i,a,t}^* = c_{i,a,t} + u_{i,a,t}^c$, where x^* denote a measured variable, x its true, unobservable value, and u the measurement error. Then rewrite equations (3) and (8) (assuming no serial correlation of the transitory shock, stationarity and omitting slope heterogeneity for simplicity) as:

$$\begin{aligned}
\Delta y_{i,a,t}^* &= \zeta_{i,a,t} + \Delta\varepsilon_{i,a,t} + \Delta u_{i,a,t}^y \\
\Delta c_{i,a,t}^* &\cong \phi\zeta_{i,a,t} + \psi\varepsilon_{i,a,t} + \Delta u_{i,a,t}^c
\end{aligned}$$

The moments of interest are:

$$\begin{aligned}
\text{var}\left(\Delta y_{a,t}^*\right) &= \sigma_\zeta^2 + 2\left(\sigma_\varepsilon^2 + \sigma_{u^y}^2\right) \\
\text{cov}\left(\Delta y_{a,t}^*, \Delta y_{a-1,t-1}^*\right) &= -\left(\sigma_\varepsilon^2 + \sigma_{u^y}^2\right) \\
\text{var}\left(\Delta c_{a,t}^*\right) &= \phi^2\sigma_\zeta^2 + \psi^2\sigma_\varepsilon^2 + 2\sigma_{u^c}^2 \\
\text{cov}\left(\Delta c_{a,t}^*, \Delta y_{a,t}^*\right) &= \phi\sigma_\zeta^2 + \psi\sigma_\varepsilon^2 \\
\text{cov}\left(\Delta c_{a,t}^*, \Delta y_{a+1,t+1}^*\right) &= -\psi\sigma_\varepsilon^2 \\
\text{cov}\left(\Delta c_{a,t}^*, \Delta c_{a-1,t-1}^*\right) &= -\sigma_{u^c}^2
\end{aligned}$$

Assume that measurement error in income is orthogonal to measurement error in consumption. Under the martingale assumption for consumption, $\sigma_{u^c}^2$ can be identified using the covariance of current and lagged consumption growth:

$$\sigma_{u^c}^2 = -\text{cov}\left(\Delta c_{a,t}^*, \Delta c_{a-1,t-1}^*\right) \tag{A2.1}$$

The variance of the permanent shock is still identified from income data:

[27] Blundell, Low and Preston (2004) contains a lengthier derivation of such an expression, including discussion of the order of magnitude of the approximation error involved.

$$\sigma_\zeta^2 = \text{var}\left(\Delta y^*_{a,t}\right) + 2\text{cov}\left(\Delta y^*_{a,t}, \Delta y^*_{a-1,t-1}\right)$$

and the partial insurance parameter ϕ from a mixture of income and consumption growth moments:

$$\phi = \frac{\text{cov}\left(\Delta c^*_{a,t}, \Delta y^*_{a,t}\right) + \text{cov}\left(\Delta c^*_{a,t}, \Delta y^*_{a+1,t+1}\right)}{\text{var}\left(\Delta y^*_{a,t}\right) + 2\text{cov}\left(\Delta y^*_{a,t}, \Delta y^*_{a-1,t-1}\right)}$$

However, σ^2_ε and $\sigma^2_{u^y}$ cannot be told apart, and ψ thus remains unidentified. It is possible however to put a lower bound on $\psi_{b,t}$ using the fact that:

$$\psi \geq \frac{\text{cov}\left(\Delta c^*_{a,t}, \Delta y^*_{a+1,t+1}\right)}{\text{cov}\left(\Delta y^*_{a,t}, \Delta y^*_{a-1,t-1}\right)}$$

Thus it is possible to argue that the estimate of ψ in Tables VIII-X is downward biased due to measurement error in income. Using estimates contained in Meghir and Pistaferri [2004], a back-of-the-envelope calculation shows that the variance of measurement error in earnings accounts for approximately 30% of the variance of the overall transitory component of earnings. Given that our estimate of ψ is close to zero in most cases, an adjustment using this inflation factor would make little difference empirically. To give an example, the estimate of ψ in Table VIII, column 1, would increase from 0.05 to 0.07.

A.3 Appendix: Estimation details

The two basic vectors of interest are:

$$\Delta \mathbf{c}_i = \begin{pmatrix} \Delta c_{i,1} \\ \Delta c_{i,2} \\ ... \\ \Delta c_{i,T} \end{pmatrix} \text{ and } \Delta \mathbf{y}_i = \begin{pmatrix} \Delta y_{i,1} \\ \Delta y_{i,2} \\ ... \\ \Delta y_{i,T} \end{pmatrix}$$

where, for simplicity, we indicate with 0 the first year in the panel (1978) and with T the last (1992), and the reference to age has been omitted. Since PSID consumption data were not collected in 1987 and 1988, the vector $\Delta \mathbf{c}_i$ is understood to have $\dim(\Delta y_i)-3$, i.e., the rows with missing consumption data have already been sweeped out from $\Delta \mathbf{c}_i$. Moreover, if the individual was not interviewed in year t, we replace the unobservable $\Delta c_{i,t}$ and $\Delta y_{i,t}$ with zeros. Conformably with the vectors above, we define:

$$\mathbf{d}^c_i = \begin{pmatrix} d^c_{i,1} \\ d^c_{i,2} \\ ... \\ d^c_{i,T} \end{pmatrix} \text{ and } \mathbf{d}^y_i = \begin{pmatrix} d^y_{i,1} \\ d^y_{i,2} \\ ... \\ d^y_{i,T} \end{pmatrix}$$

where $d^c_{i,t} = 1\left\{\Delta c_{i,t} \text{ is not missing}\right\}$ and $d^y_{i,t} = 1\left\{\Delta y_{i,t} \text{ is not missing}\right\}$. Overall, this notation allows us to handle in a simple manner the problems of unbalanced panel data and of missing consumption data in 1987 and 1988.

Stacking observations on Δy and Δc (and on d^c and d^y) for each individual we obtain the vectors:

$$\mathbf{x}_i = \begin{pmatrix} \Delta \mathbf{c}_i \\ \Delta \mathbf{y}_i \end{pmatrix} \text{ and } \mathbf{d}_i = \begin{pmatrix} \mathbf{d}^c_i \\ \mathbf{d}^y_i \end{pmatrix}$$

Now we can derive:

$$\mathbf{m} = vech\left\{\left(\sum_{i=1}^{N} \mathbf{x}_i \mathbf{x}_i'\right) \oslash \left(\sum_{i=1}^{N} \mathbf{d}_i \mathbf{d}_i'\right)\right\}$$

where $\oslash$ denotes an elementwise division. The vector $\mathbf{m}$ contains the estimates of $cov\left(\Delta y_t, \Delta y_{t+s}\right)$, $cov\left(\Delta y_t, \Delta c_{t+s}\right)$, and $cov\left(\Delta c_t, \Delta c_{t+s}\right)$, a total of $T\left(2T+1\right)$ unique moments).[28] To obtain the variance-covariance matrix of $\mathbf{m}$, define conformably with $\mathbf{m}$ the individual vector:

$$\mathbf{m}_i = vech\left\{\mathbf{x}_i\mathbf{x}_i'\right\}$$

The variance-covariance matrix of $\mathbf{m}$ that can be used for inference is:

$$\mathbf{V} = \left[\sum_{i=1}^{N}\left(\left(\mathbf{m}_i - \mathbf{m}\right)\left(\mathbf{m}_i - \mathbf{m}\right)'\right) \circledast \mathbf{D}_i\right] \oslash \left(\sum_{i=1}^{N}\mathbf{D}_i\right)$$

where $\mathbf{D}_i = vech\left\{\mathbf{d}_i\mathbf{d}_i'\right\}$ and $\circledast$ denotes an elementwise product. The square roots of the elements in the main diagonal of $\mathbf{V}$ provide the standard errors of the corresponding elements in $\mathbf{m}$.

What we do in the empirical analysis is to estimate models for $\mathbf{m}$:

$$\mathbf{m} = f\left(\mathbf{\Lambda}\right) + \mathbf{\Upsilon}$$

where $\mathbf{\Upsilon}$ captures sampling variability and $\mathbf{\Lambda}$ is the vector of parameters we are interested in (the variances of the permanent shock and the transitory shock, the partial insurance parameters, etc.). For instance the mapping from $\mathbf{m}$ to $f\left(\mathbf{\Lambda}\right)$ is:

$$\begin{pmatrix} \text{var}\left(\Delta c_1\right) \\ \text{cov}\left(\Delta c_1, \Delta c_2\right) \\ ... \\ \text{cov}\left(\Delta c_1, \Delta c_T\right) \\ ... \end{pmatrix} = \begin{pmatrix} \phi^2\text{var}\left(\zeta_1\right) + \psi^2\text{var}\left(\varepsilon_1\right) + \text{var}\left(\xi_1\right) + 2\sigma^2_{u^c} \\ -\sigma^2_{u^c} \\ ... \\ 0 \\ ... \end{pmatrix} + \mathbf{\Upsilon}$$

We solve the problem of estimating $\mathbf{\Lambda}$ by minimizing:

$$\min_{\mathbf{\Lambda}}\left(\mathbf{m} - f\left(\mathbf{\Lambda}\right)\right)'\mathbf{A}\left(\mathbf{m} - f\left(\mathbf{\Lambda}\right)\right)$$

where $\mathbf{A}$ is a weighting matrix. Optimal minimum distance (OMD) imposes $\mathbf{A} = \mathbf{V}^{-1}$, equally weighted minimum distance (EWMD) imposes $\mathbf{A} = \mathbf{I}$, and diagonally-weighted minimum distance (DWMD) requires that $\mathbf{A}$ is a diagonal matrix with the elements in the main diagonal given by $diag\left(\mathbf{V}^{-1}\right)$.

For inference purposes we require the computation of standard errors. Chamberlain [1984] shows that these can be obtained as:

$$\widehat{var\left(\widehat{\mathbf{\Lambda}}\right)} = \left(\mathbf{G}'\mathbf{A}\mathbf{G}\right)^{-1}\mathbf{G}'\mathbf{A}\mathbf{V}\mathbf{A}\mathbf{G}\left(\mathbf{G}'\mathbf{A}\mathbf{G}\right)^{-1}$$

where $\mathbf{G} = \frac{\partial f(\mathbf{\Lambda})}{\partial \mathbf{\Lambda}}\Big|_{\mathbf{\Lambda}=\widehat{\mathbf{\Lambda}}}$ is the Jacobian matrix evaluated at the estimated parameters $\widehat{\mathbf{\Lambda}}$.

[28] In practice there are less than $T\left(2T+1\right)$ moments because data on consumption are not available all years.