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FINANCIAL INTEGRATION AND CONSUMPTION SMOOTHING*

Tullio Jappelli and Luigi Pistaferri

We test if financial integration improves household consumption smoothing using microeconomic data. We find that the process of financial market integration and liberalisation brought about by the introduction of the euro has not affected the sensitivity of consumption with respect to income shocks in Italy. This article also makes a significant contribution from a methodological point of view, because our procedure does not require that consumption and income are available in the same panel data set. It can therefore be applied in all countries in which repeated cross-sectional consumption data can be combined with panel data on income.

The European Monetary Union (EMU) has removed exchange rate risk and lowered cross-border transaction costs, opening the possibility for the creation of a fully integrated continental financial market comparable with that of the US. To what extent has this process of regulatory reform affected the ability of households to diversify, insure and shoulder risks? This article attempts to answer this question, which is at the heart of the burgeoning literature on the links between regulation, finance and real economic activity. To answer the question, we study the effect of the euro policy shift with Italian household-level income and consumption data spanning two decades (1987–2006).

Models of intertemporal choice imply that consumers use credit and insurance markets to smooth, at least in part, income shocks. This fundamental implication of the theory suggests that consumption should be less sensitive to income shocks after a period of financial market liberalisation. Applying Cochrane (1991) and Mace (1991) seminal contributions to aggregate data, Sørensen *et al.* (2007) test if the response of country consumption growth to country idiosyncratic income shocks falls after the introduction of the euro. These tests rely on the strong assumption that countries are populated by identical consumers. Furthermore, it is by now well established in the microeconometric literature that the bulk of income variability is because of individual-specific shocks, rather than to region or countrywide shocks.

In this article, we develop a new empirical strategy for testing if financial liberalisation improves consumers' ability to hedge against income shocks. Our analysis is performed at the cohort level using Italian data from the 1987–2006 Survey of Household Income and Wealth (SHIW). We build on work by Deaton and Paxson (1994), Blundell and Preston (1998) and Blundell *et al.* (2008), and decompose the change in the variance of consumption into a component that depends on the variance of permanent

* Corresponding author: Tullio Jappelli, Department of Economics, University of Naples Federico II, Via Cinthia 45, 80146 Napoli, Italy. Email: tullio.jappelli@unina.it.

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income shocks and one that depends on the change in the variance of transitory shocks. We then test if the process of financial market integration and liberalisation brought about by the introduction of the euro has affected the sensitivity of consumption to income shocks. The test allows us to recover two key policy parameters – the sensitivities of consumption to permanent and transitory income shocks – and to check if the two parameters have changed after the introduction of the euro.

This article makes a contribution also from a methodological point of view. We use panel data on income to identify non-parametrically a time series of the variances of the income shocks for each cohort. However, we rely on repeated cross-sectional data to construct the cohort-specific variances of income and consumption. We then combine panel data and repeated cross-sectional data for each cohort to identify the sensitivity of consumption with respect to income shocks and to test if it has declined after the introduction of the euro. Our procedure does not require that consumption and income are available in the same panel data set, whereas Blundell *et al.* (2008) requires that panel data on income and consumption are available for the same households and for a long number of years. Our procedure can therefore be applied to situations in which there are repeated cross-sections containing data on consumption and income but panel data exist only for income.

The results indicate that the sensitivity of consumption to income shocks tends to decline after the introduction of the euro but such effect is not statistically different from zero. We check that this result applies to different definitions of income, using alternative measures of financial integration and adding the restrictions that the theory imposes on the joint behaviour of consumption and income. Finally, to distinguish the EMU effect from potential confounds more sharply, we compare statistically the dynamics of consumption and income inequality in Italy (a country that joined EMU) with the dynamics of the same variables in the UK (a country that did not join). For this robustness check, we draw consumption and income data from the UK Family Expenditure Survey (FES) and longitudinal income data from the British Households Panel Survey (BHPS), using the same methodology and specification as in the Italian case. The use of UK data also demonstrates the advantage of our methodological point (FES is a pure repeated cross-section with consumption and income data, whereas BHPS is a panel of income but not consumption).

This article is organised as follows. Section 1 reviews the literature on the effect of financial market integration on risk-sharing opportunities and consumption smoothing. Section 2 discusses the macroeconomic developments in the euro zone and Italy before and after the introduction of the euro. Section 3 explains how changes in the variance of consumption over time can signal changes in consumption smoothing. Section 4 presents the Italian data and explains how we construct the three ingredients of our test: consumption inequality income inequality and the variance of the income shocks. Section 5 presents the baseline results and Section 6 the results of the robustness analysis. Section 7 concludes.

1. Financial Market Integration, Risk Sharing and Consumption Smoothing

Economic theory predicts that the process of financial market integration should facilitate consumption smoothing and risk-sharing opportunities. First of all, it should allow households to hold more diversified equity portfolios and, in particular, to diversify the portion of risk that arises from country-specific shocks. But most importantly, integration should spur the efficiency of financial intermediaries and markets in countries where the financial system is more backward and more heavily regulated, fostering the growth of domestic financial markets and the entry of foreign banks, and improving access to credit for households. As a result, country-specific shocks should have a smaller effect on consumption when international financial markets are integrated, since they can be diversified away by borrowing abroad or holding foreign assets. At the same time, easier access to credit should help domestic borrowers to buffer specific shocks to their incomes.[1]

Accordingly, a whole line of research studies the covariance of consumption across different regions or countries to test if financial markets afford full risk sharing to consumers located in different jurisdictions. Conditional on consumers exploiting all risk-sharing opportunities, consumption growth of all regions or countries should be perfectly correlated when financial markets are integrated and depend only on common (non-diversifiable) shocks. This important point has been initially recognised and applied to US microeconomic data by Cochrane (1991) and Mace (1991) and later brought to bear on macroeconomic data by Obstfeld (1994, ch. 2), Wincoop (1994) and Townsend (1994), among others.

The risk-sharing approach is also capable of distinguishing the contribution of different financial markets and public tax-transfer mechanisms. Using US data for 1963–90, Asdrubali *et al.* (1996) develop an accounting framework to decompose the cross-sectional variance of individual states' gross output. They identify three channels through which risk sharing can occur. First, in a monetary union risk can be shared through cross-ownership of real and financial assets and thus people can smooth their income stream relative to their output stream. Second, the federal government can insure some of the income variability through taxes and transfers, thereby creating a wedge between income earned and after-tax income. Third, people could smooth consumption by owning a diversified asset portfolio and undertaking intertemporal borrowing and lending. Applying such framework, Asdrubali *et al.* (1996) find that in the US, 39% of the shocks are absorbed via capital market smoothing, 13% via the fiscal channel and 23% via the credit market, whereas the remaining 25% are not smoothed. Sørensen and Yosha (1998) and Kalemli-Ozcan *et al.* (2003, 2005) apply the same approach to the EU and the OECD for the time interval 1966–90. They find that the unsmoothed residual, estimated to be around 60%, is much larger than in the US. They also report that one half of the smoothed income risk is achieved by national government budget deficits and the other half by corporate savings. Roubini *et al.* (2007) extend the analysis to 2006, and find that risk sharing in the EMU is still significantly lower than in the US but that it has significantly improved over time in the euro zone and during the EMU period.[2]

[1] See Jappelli and Pagano (2010) for a survey of the real effects of financial market integration in the context of the EMU.

[2] Sørensen *et al.* (2007) also report that there has been an increase in risk sharing among OECD countries between 1993 and 2003. They document that this increase is correlated with the concomitant reduction in home bias, especially for equities but this finding is weaker for EU countries.

These results are informative about risk sharing across countries or regions but not about risk sharing within a country. In fact, risk-sharing tests using aggregate data assume that there is a representative agent within each country (or region of a country), implicitly assuming that agents are fully insured against person-specific shocks (such as unemployment, low productivity due to health shocks and disability). In this article, we fill a gap in the literature and provide a test for the effect of financial integration on risk sharing and consumption smoothing opportunities based on microeconomic data. We assess with household-level data if the response to income shocks has changed in Italy after the 1999 introduction of the euro. One advantage of using household-level data is that the structure of the financial system can be considered exogenous with respect to the choice of individual consumers. Using microeconomic data, one can also tackle issues that cannot be addressed with country-level data. For instance, we are able to test whether financial integration affects disproportionately some groups of households, such as specific cohorts or education groups.

While no previous study has analysed the impact of financial integration on consumption using household-level data, empirical evidence with firm-level data exists.[3] Alfaro and Charlton (2007) show that reducing restrictions on international capital flows enhances firm entry and other measures of entrepreneurship. Bertrand *et al.* (2007) find that following the banking deregulation carried out by the French Banking Act of 1985, banks became less willing to bail out poorly performing firms, whereas firms in more bank-dependent sectors became more likely to undertake restructuring activities.[4]

2. The Process of Financial Market Integration

The introduction of the euro has eliminated exchange rate risk, as well as the costs arising from exchange rate transactions within the euro zone. Therefore, it has directly removed one of the main barriers to financial integration. The process leading to monetary unification also triggered a sequence of policy actions and private sector responses that swept aside many other regulatory barriers to financial integration: controls on capital flows were removed, banking and financial service directives were passed to create a level playing field in credit and securities market, and the rules governing the issuance of public debt were homogenised. In short, the EMU has been the single most important policy-induced innovation in the international financial system since the collapse of the Bretton-Woods system, opening the possibility for the creation of a fully integrated European financial market comparable to that of the US. Jappelli and Pagano (forthcoming) describe these developments, and the effect of the EMU on financial market integration, investment, growth, ability to respond to macroeconomic shocks and risk-sharing opportunities.

[3] Several studies using firm-level data document that financial development has a positive effect on access to finance and entry of new firms (Guiso *et al.*, 2004*a*; Aghion *et al.*, 2007; Klapper *et al.*, 2004).

[4] Recent microeconomic evidence also throws light on the role that international financial integration can play in improving the allocation of capital across firms. Galindo *et al.* (2007) use firm-level panel data from 12 Latin American countries to investigate whether capital account liberalisation has increased the share of investment going to firms with a higher marginal return to capital.

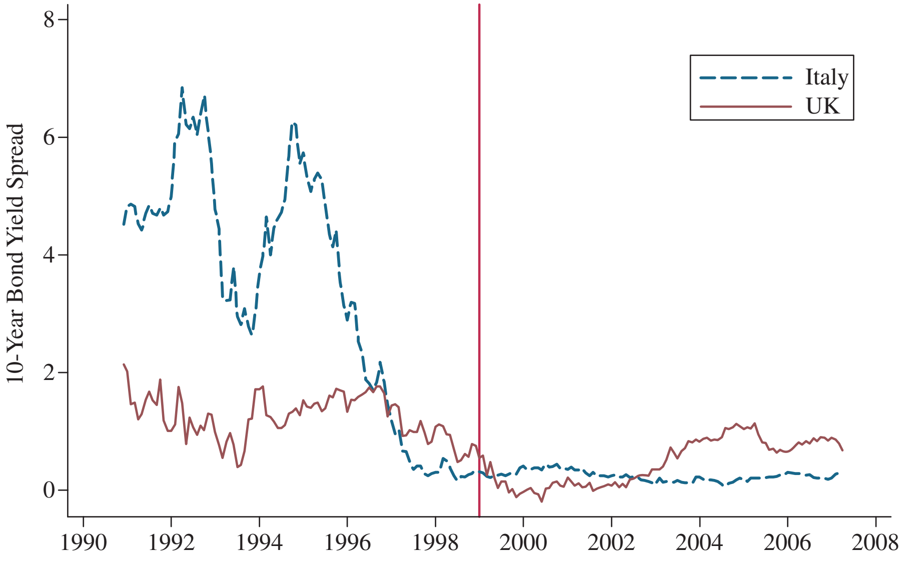

Fig. 1. *Ten-Year Benchmark Bond Yield Spread in Italy and the UK*
Note. Yield differentials are computed as the difference relative to the yield on German 10-year benchmark bonds, based on monthly data (end-of-month observations).
Source. Datastream.

The combined effect of EMU and concomitant institutional changes translated into a convergence of interest rates on the eve of monetary unification. Inter-bank money markets in the euro area fully integrated, whereas other interest rates have converged rapidly. In Italy, the convergence towards zero of the spread over the German yield is dramatic, as shown in Figure 1.[5] The Figure also plots the spread for the UK, a country that is part of the European Union but did not join the Euro in 1999. As shown in the Figure, the UK featured a lower spread before 1999. However, after the introduction of the euro, the Italian spread has been on average smaller and less volatile than in the UK, which was pursuing an independent monetary policy. The UK therefore represents an interesting country to compare with Italy, an issue that we will take up econometrically in Section 6.[6]

In Italy, the most important development of financial market integration affecting consumers is the growth of the consumer credit and mortgage markets, the two financial markets that are more directly related to households' ability to smooth income fluctuations. Historically, the Italian mortgage and consumer credit markets were severely limited by regulation, judicial inefficiency and high enforcement costs. Chiuri and Jappelli (2003) document that the cost of mortgage foreclosure, the length of trials and judicial

[5] The convergence was similar in other non-core EMU participants: Finland, Ireland, Portugal and Spain, and later Greece, which joined the euro area at the beginning of 2001.

[6] Financial integration in other markets has proceeded more slowly. Integration of equity markets has been less pronounced, reflecting obstacles to cross-border trading and different national company laws. Nevertheless, the share of equity held in other euro area countries increased significantly between 1999 and 2007, reaching almost 30% (European Commission, 2008). In the banking sector, the initial wave of consolidation in the euro-area occurred almost exclusively within national borders, and cross-border retail banking remains rather limited within the euro area.

inefficiency in Italy are higher than in countries at a similar level of financial development.[7] Casolaro *et al.* (2006, ch. 4) also stress that, compared with other countries, Italy features a lower level of social capital and trust, which affects real and financial transactions.

Despite the fact that the Italian mortgage and consumer credit markets are still small by international standards, the process of European financial integration and the associated fall in interest rates has increased considerably households' incentives to borrow. Furthermore, financial integration has spurred increasing competitive pressure, reducing the cost of debt and increasing the supply of loans. This is documented in Figure 2, which shows that the household debt–gross domestic product (GDP) ratio more than tripled from 9% in 1986 to almost 30% in 2006, with particularly strong growth around 1999. National regulatory changes also played an important role, with the removal of regulations on entry, limitations of geographical span of lending and separation of long- and short-term lending. Specific mortgage regulation has also eased considerably, and loan maturities and loan-to-value ratios have gradually increased. The development of a credit reporting system and credit scoring techniques in the mid-1990s has improved the quality of information on prospective borrowers, benefiting the performance of household debt markets. In short, even though the household debt market still lags behind other industrialised nations, the market has grown at double digit rates, especially around and after the 1999 introduction of the euro. In the next Section, we show how we will use the euro policy shift to identify the potential effect of financial market integration on consumption.

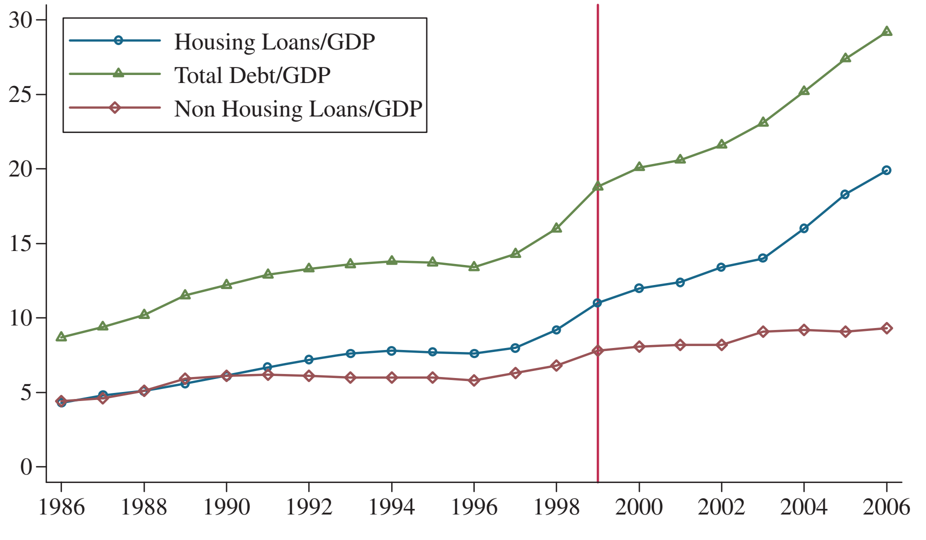

Fig. 2. *Household Debt–GDP in Italy*

Sources. Bank of Italy Annual Report, Statistical Appendix, various years. Consumer Credit and Lending to Households in Europe – ECRI 2006 Statistical Package. Bruxelles: European Credit Research Institute.

[7] A further reason for the relatively thin mortgage and consumer credit markets is the presence of informal arrangements and various forms of intergenerational transfers (bequests, *inter vivos* transfers, help for down payment or outright purchase, free housing or co-residency), partly overcoming borrowing constraints and reducing the need for mortgage credit.

3. The Empirical Strategy

We rely on the covariance restriction implied by the permanent income hypothesis (PIH) to check if the variance of consumption tracks less closely the variance of income after the introduction of the euro. For this purpose, we rely on standard assumptions about the evolution of household income (Deaton, 1991; Carroll, 1997; Blundell and Preston, 1998). In particular, we define income as the sum of after-tax family earnings and transfers (we thus exclude income from assets), and decompose it into three parts: a deterministic component, a permanent component and a transitory shock (assuming for the time being that income is measured without error):

$$\ln y_{i,a,t} = \boldsymbol{x}'_{i,a,t}\boldsymbol{\beta} + P_{i,a,t} + e_{i,a,t}, \tag{1}$$

where

$$P_{i,a,t} = P_{i,a-1,t-1} + u_{i,a,t},$$

and i, a and t are subscripts for individual, age and time, respectively.

The decomposition of income shocks into transitory and permanent components dates back to Friedman (1957). Some of the income shocks are transitory (mean reverting) and their effect does not last long. Examples include fluctuations in overtime labour supply, bonuses, lottery prizes and bequests. However, some of the innovations to earnings are highly persistent (non-mean reverting) and their effect accumulates over time. Examples of permanent innovations are generally associated with job mobility, promotions, lay-offs and severe health shocks. We comment on the plausibility of this income process in Section 4.2.

In this article, we study the effect of financial integration on consumption smoothing, not on income smoothing. We therefore assume that income evolves exogenously, and that it is the only source of idiosyncratic risk faced by consumers.[8] Recent papers have explored the implications of endogenising income for consumption allocations through human capital accumulation, job search, labour supply and cross-firm mobility. Accounting for endogenous income, however, is beyond the purpose of this study.[9]

Assume that individuals of all cohorts enter the labour market at age a_0. For an individual aged a in year t (and hence born in year $b = t - a$), we have:

$$\ln y_{i,a,t} = \boldsymbol{x}'_{i,a,t}\boldsymbol{\beta} + \pi_{i,a_0,t-a+a_0} + \sum_{j=a_0+1}^{a} u_{i,j,t-a+j} + e_{i,a,t}.$$

[8] Our test is designed to estimate the sensitivity of consumption with respect to transitory and permanent shocks to after-tax income (i.e. after the smoothing by employers and/or government has taken place). It is indeed an interesting question whether there is any income smoothing provided by employers and government. Unfortunately, studying this issue empirically is difficult. For example, the variance of income that we measure may have already been smoothed by, say, implicit contracts with the firm. To check whether this is true and whether smoothing has changed over time would require data on workers' productivity (rather than just wages), which are seldom available in standard data sets. In the robustness analysis, we check whether our test for financial integration is sensitive to using a measure of income that excludes transfers (public and private).

[9] See Low *et al.* (2010), Heathcote *et al.* (2010) and Postel-Vinay and Thuron (2010) for recent applications.

The term π is the initial draw of the permanent component P. It represents differences in initial abilities and other fixed characteristics among individuals entering the labour market in the same year, that is, individuals of the same cohort. We take the variance of the income process with respect to all individuals of the same cohort, so that:

$$\mathrm{var}_b(\ln y_{i,a,t}) = \mathrm{var}_b(\pi_{i,a_0,t-a+a_0}) + \sum_{j=a_0+1}^{t} \mathrm{var}_b(u_{i,j,t-a+j}) + \mathrm{var}_b(e_{i,a,t}), \tag{2}$$

where $\mathrm{var}_b(\cdot)$ denotes the variance for cohort born in year b. For convenience, we have omitted the contribution of the observable characteristics x, which do not play any role for describing the evolution of income inequality. Equation (2) indicates that the variance of income of each cohort in a given year is the sum of the variance of initial conditions, the cumulative variances of permanent shocks and the variance of the transitory shocks in that year. Note that we have made the assumption that the three stochastic components π, u and e are mutually uncorrelated at all lags. We also assume that u and e are not serially correlated. However, we do not need to assume that u and e are covariance-stationary.

Following Blundell *et al.* (2008), we obtain a similar decomposition of the variance of consumption, starting from an approximation of the Euler equation. To derive such approximation, let us assume that consumers have constant relative risk aversion preferences, that income follows the process (1), and that it is the only source of uncertainty. Individual consumption can then be written as:

$$\ln c_{i,a,t} \approx \ln c_{i,a-1,t-1} + z'_{i,a,t}\gamma + \phi u_{i,a,t} + \psi e_{i,a,t}. \tag{3}$$

Equation (3) nests many consumption models. It shows that consumption growth depends on preference shifts z (such as age and family size) and income shocks. The response of consumption to income shocks is captured by the parameters ϕ and ψ.[10] As discussed by Jappelli and Pistaferri (2010), these parameters depend on the planning horizon, interest rates, preference parameters and the structure of credit and insurance markets. According to the permanent income model, consumption responds fully to permanent income shocks ($\phi \approx 1$), whereas transitory shocks have negligible effects ($\psi \approx 0$) because consumers use assets to smooth temporary income fluctuations. The buffer stock model delivers similar implications.[11] If there are complete markets, individual consumption is completely insulated from transitory as well as permanent

[10] Unlike risk-sharing regressions *à la* Mace (1991), our approach allows estimating the marginal propensity to consume with respect to shocks of various nature and persistence simultaneously, rather than to specific episodes (like weather fluctuations or job loss). The drawbacks are that it assumes that income and consumption follow a particular process, and it is more demanding in terms of data.

[11] Simulation results produced by Carroll (2009) show that with constant relative risk aversion, impatient consumers and an income process similar to the one we use, the implication of the PIH that transitory income shocks have a negligible impact on consumption still holds true. Permanent shocks, however, have a somewhat lower impact in buffer stock models. In fact, in such models permanent income shocks reduce the ratio of wealth to permanent income, thus increasing also precautionary saving. Under a wide range of parameter values, Carroll shows that in this class of models the marginal propensity to consume out of a permanent income shock is about 0.9. Kaplan and Violante (2010) find qualitatively similar results in simulations of the Bewley model. In particular, they find that, on average, the sensitivity of consumption with respect to permanent income shocks is 0.77.

shocks ($\phi = \psi = 0$). Finally, models with partial insurance predict that consumers are able to also insure permanent shocks to a larger extent than in the PIH ($\psi \approx 0$ and $0 < \phi < 1$).

The parameter ψ in (3) represents the extent to which consumption responds to income over and above the amount warranted by the PIH, that is, the excess sensitivity of consumption to transitory income shocks. Some authors rationalise excess sensitivity by appealing to the presence of binding liquidity constraints in each period. Laibson (1997) shows that it is the equilibrium outcome for consumers with hyperbolic preferences.[12] Others term it rule-of-thumb model to indicate a situation in which consumption tracks income closely, even when individuals have accumulated assets in previous periods. The model is an interesting case to study because it approximates the behaviour of consumers with short horizons, limited resources or hyperbolic discount factors, giving an upper bound for the sensitivity of consumption to income shocks.[13]

Consistent with the models' predictions, we denote ψ as the *excess sensitivity parameter*: lower values of this parameter imply that consumers are more able to smooth transitory income shocks by borrowing and lending. We also denote ϕ as the *insurance parameter*: lower values of this coefficient signal that consumers have access to more insurance opportunities, and therefore there is less tracking of consumption to permanent income shocks. As we shall see, in the empirical analysis, we shall allow the two parameters to vary over time to capture changes in the degree of consumption smoothing.

One would expect that the process of financial market integration and the associated credit market development and consumption smoothing opportunities translate in to a reduction over time in the sensitivity of consumption to transitory shocks (ψ). The effect of financial market integration on the sensitivity to permanent shocks (ϕ) is less clear-cut. On the one hand, insurance opportunities increase with financial market integration, as consumers can more easily diversify risk by holding foreign assets. But financial integration may also diminish the role of fiscal policy in countries with initially less developed financial markets (Bertola, 2007, ch. 6).[14] For these reasons, one should expect that financial integration might impact consumption primarily through a change in ψ rather than in ϕ.

As in the case of income, for an individual aged a in year t who enters the labour market at age a_0, we can rewrite (3) as:

[12] In the hyperbolic consumer model, individuals have preferences that change over time (there are different selves in different periods). In the model proposed by Laibson (1997), self $t - 1$ chooses assets a_{t-1} to constrain the consumption of self t. This is done by keeping most assets invested in an illiquid instrument. Hence, at any point in time, the consumer is effectively liquidity constrained, even though the constraint is self-imposed. Laibson (1997) shows that in equilibrium, consumption is exactly equal to the current level of cash flow, or total income.

[13] An upper bound for the excess sensitivity parameter is $\psi = 1$. This case can arise if consumers are myopic and set consumption equal to income ($\ln c_{a,i,t} = \ln y_{a,i,t}$), so that consumption responds fully to permanent and transitory income shocks. This model has been often proposed as a simple, yet extreme alternative to the PIH to describe the behaviour of households that do not use savings to buffer income shocks but spend all that they receive.

[14] Financial development may lower their need for government-provided insurance, insofar as the markets will be able to provide the risk-sharing services that people would otherwise expect from the social security system and the welfare state. This would allow these countries to focus their social welfare systems more closely on its redistributive role, and away from risk-sharing.

$$\ln c_{i,a,t} = \ln c_{i,a_0,t-a+a_0} + \phi \sum_{j=a_0+1}^{a} u_{i,j,t-a+j} + \psi \sum_{j=a_0+1}^{a} e_{i,j,t-a+j},$$

where $c_{i,a_0,t-a+a_0}$ reflects initial differences in preferences and endowments of individuals that belong to cohort *b*. For convenience, we have omitted the contribution of the observable characteristics *z*, which do not play any role for describing the evolution of consumption inequality. Taking the variance of consumption for these individuals, we obtain:

$$\text{var}_b(\ln c_{i,a,t}) = \text{var}(\ln c_{i,a_0,t-a+a_0}) + \phi^2 \sum_{j=a_0+1}^{a} \text{var}_b(u_{i,j,t-a+j}) + \psi^2 \sum_{j=a_0+1}^{a} \text{var}_b(e_{i,j,t-a+j}). \quad (4)$$

Equation (4) indicates that the variance of consumption of each cohort in year *t* is the sum of the variance of initial conditions and of the cumulative variances of permanent and transitory shocks until year *t*, weighted by the square of the insurance and excess sensitivity parameters, respectively.

Consider now the changes in the cross-sectional income and consumption variances, that is, the first difference of (2) and (4):

$$\Delta\text{var}(\ln y_{i,a,t}) = \text{var}(u_{i,a,t}) + \Delta\text{var}(e_{i,a,t}), \quad (5)$$

$$\Delta\text{var}(\ln c_{i,a,t}) = \phi^2\text{var}(u_{i,a,t}) + \psi^2\text{var}(e_{i,a,t}), \quad (6)$$

where we have omitted the subscript *b* on the variance terms to simplify notation.

Taking the first differences has two advantages. First, it removes the 'fixed' effects (the initial conditions) that are specific to each cohort. Second, because of the martingale structure of the errors, the first difference operator also removes the 'history' of inequality induced by permanent shocks, and hence imposes less strict data requirements.

In (5), the change in income inequality from one year to the next for a given cohort is because of the arrival of permanent and transitory shocks. In the absence of transitory shocks, income inequality unambiguously rises because of the spreading out effect induced by permanent shocks. Income inequality is also affected by the change in the variance of transitory shocks, so overall income inequality may fall if the inequality component induced by transitory shocks declines over time and the variance of permanent shocks is small.

Equation (6) highlights the determinants of changes in consumption inequality, and the second column of Table 1 illustrates the implications of various models of consumption behaviour for such changes. In the PIH ($\phi = 1$, $\psi = 0$), consumption inequality spreads out over time, an implication of the model first pointed out by Deaton and Paxson (1994). In this model, only the presence of non-stationary measurement error in consumption may explain a possible fall in consumption inequality. In models where there is excess sensitivity of consumption to transitory income shocks ($\psi > 0$), the change in the variance of consumption within each cohort reflects also the variance of transitory shocks. Models with partial insurance ($0 < \phi < 1, \psi = 0$) also predict a fanning out of cohort inequality, albeit at a slower pace than in the PIH.

Table 1

Implications of Various Models for Consumption and Income Inequality

Model	Change in the variance of consumption = $\Delta\text{var}\ln(c_{i,a,t})$	Difference-in-difference between change in the variance of income and variance of consumption = $\Delta\text{var}\ln(y_{i,a,t}) - \Delta\text{var}\ln(c_{i,a,t})$	Restrictions
PIH	$\text{var}(u_{i,a,t})$	$\Delta\text{var}(e_{i,a,t})$	$\phi = 1,\ \psi = 0$
Partial insurance	$\phi^2\text{var}(u_{i,a,t})$	$(1-\phi^2)\text{var}(u_{i,a,t}) + \Delta\text{var}(e_{i,a,t})$	$0 < \phi < 1,\ \psi = 0$
Complete markets	0	$\text{var}(u_{i,a,t}) + \Delta\text{var}(e_{i,a,t})$	$\phi = \psi = 0$
Excess sensitivity	$\text{var}(u_{i,a,t}) + \psi^2\text{var}(e_{i,a,t})$	$(1-\psi^2)\text{var}(e_{i,a,t}) - \text{var}(e_{i,a,t-1})$	$\phi = 1,\ 0 < \psi \leq 1$

Under complete markets ($\phi = \psi = 0$), consumers are insulated from all shocks, and cohort consumption inequality is constant over time.

Our empirical specification will consist of estimating the determinants of the *divergence* between changes in income and consumption inequality, that is, the difference-in-difference of (5) and (6):

$$\Delta\text{var}(\ln y_{i,a,t}) - \Delta\text{var}(\ln c_{i,a,t}) = (1-\phi^2)\text{var}(u_{i,a,t}) + (1-\psi^2)\text{var}(e_{i,a,t}) - \text{var}(e_{i,a,t-1}). \quad (7)$$

As we shall see, in the data, we observe periods in which income and consumption inequality exhibit different trends, and (7) can be used to understand the forces behind this divergence. The third column of Table 1 reports the implications of the various models of consumption for the difference-in-difference between the change in income and consumption inequality.

The impact of the variance of permanent shocks depends on the particular consumption model considered: the impact is one-for-one in the complete market model (because here the change in the variance of consumption is zero), positive but less then one in the partial insurance case and zero in the PIH. With the exception of models with excess sensitivity, the change in the variance of transitory shocks impacts one-for-one on the difference between the changes in the income and consumption variances.[15] Thus, in models in which households smooth transitory income shocks, one needs an increase in income instability, that is, $\Delta\text{var}(e_{i,a,t}) > 0$, to generate a divergence between income and consumption inequalities, regardless of trends in the variance of the permanent shocks or changes in the degree of insurance. The complete markets case can also generate a diverging path but because consumption inequality does not grow, it predicts that the divergence equals the growth in income inequality, a restriction that can be easily tested.

In this article, we are particularly interested in estimating the impact of EMU on the excess sensitivity and insurance coefficients. We therefore define a dummy indicator E for the post-1999 observations and rewrite (7) allowing the sensitivity of

[15] Note that if consumers are myopic and set consumption equal to income, the variance of consumption tracks the variance of income, $\text{var}(\ln c_{i,a,t}) = \text{var}(\ln y_{i,a,t}) = \text{var}(u_{i,a,t}) + \Delta\text{var}(e_{i,a,t})$, so that the difference between the two is zero.

consumption to income shocks to be affected linearly by the introduction of the euro:

$$\begin{aligned}\Delta\mathrm{var}(\ln y_{i,a,t}) - \Delta\mathrm{var}(\ln c_{i,a,t}) = {} & [1-(\phi+\phi_E E)^2]\mathrm{var}(u_{i,a,t}) \\ & + [1-(\psi+\psi_E E)^2]\mathrm{var}(e_{i,a,t}) - \mathrm{var}(e_{i,a,t-1}).\end{aligned} \tag{8}$$

We test for the effect of financial integration by looking at the p-value of the joint null hypothesis of no EMU effect ($\varphi_E = \psi_E = 0$), against the hypothesis that these parameters are negative, either because households are more able to insure permanent shocks or because they are more able to smooth transitory shocks. Note that the change in the consumption response to shocks after EMU might reflect both changes in insurance and credit market smoothing opportunities, as well as the reduction in interest rates associated with the introduction of the euro. In the robustness analysis, we study if specific measures of financial integration, such as the interest rate spread, give different estimates of the coefficients of interest. As explained earlier, the effect of financial integration on the ability to insure permanent shocks is not as clear-cut as for transitory shocks. Since we expect financial market integration to affect primarily the sensitivity of consumption to transitory shocks, in some specifications we impose the restriction $\varphi_E = 0$ and test only that sensitivity of consumption to transitory income shocks has not changed over time.

We run the regression based on (8) using two kinds of data: repeated cross-sections on income and consumption, and panel data on income. Repeated cross-sections on income and consumption allow us to identify the changes in cross-sectional income and consumption inequality, that is, the left-hand side of (8). Panel data on income allow us to identify the variances of income shocks, that is, the right-hand side variables of (8). Omitting for simplicity, the contribution of the observable characteristics x, as in Meghir and Pistaferri (2004), we identify the cohort-specific variances of income shocks non-parametrically using:

$$\mathrm{E}[(\ln y_{i,a,t} - \ln y_{i,a-1,t-1})(\ln y_{i,a+1,t+1} - \ln y_{i,a-2,t-2})] = \mathrm{var}(u_{i,a,t}), \tag{9}$$

$$-\mathrm{E}[(\ln y_{i,a,t} - \ln y_{i,a-1,t-1})(\ln y_{i,a+1,t+1} - \ln y_{i,a,t})] = \mathrm{var}(e_{i,a,t}). \tag{10}$$

Note that estimation of (9) requires four years of data on each household (from $t-2$ to $t+1$), while estimation of (10) requires three years of data (from $t-1$ to $t+1$). Furthermore, if income is measured with classical i.i.d. error, one can prove that (9) still identifies the variance of permanent shocks, while (10) will identify the sum of the variance of transitory shocks and the variance of the measurement error. Assuming consumption is independent of measurement error in income, one can further prove that the estimate of ϕ is unbiased, while the estimate of ψ is upward biased.[16] Clearly, unless measurement error changes systematically over time, there is no reason to believe that our test of financial integration is affected by

[16] In particular, $\mathrm{p\,lim}\,\hat{\psi} = \sqrt{1-\lambda(1-\psi^2)}$, where λ is the signal-to-noise ratio, that is, the ratio of the variance of transitory shocks to the sum of the variance of transitory shocks and the variance of the measurement error.

measurement error in income. As for measurement error in consumption, as long as it is an i.i.d. classical error, its variance will vanish when taking first differences of consumption variances (6).[17]

From a methodological point of view, the test can be applied to situations in which income and consumption are not available in the same dataset, or perhaps more usefully, to situations in which there are repeated cross-sections on consumption and income, but panel data exist for income but not for consumption. Examples of applicability include the US, where the Consumer Expenditure Survey (CEX) provides repeated cross-sectional data on consumption and the Panel Study of Income Dynamics (PSID) provides panel data on income. As we shall see in Section 6, another application is the UK, where repeated cross-section data on consumption and income data are available from the FES and panel data on income from the BHPS.

4. The Data

Our test provides the first attempt to evaluate the impact of financial market integration on consumption using household-level data, requiring panel data on income to estimate the cohort variances of transitory and income shocks, and repeated cross-sectional data on consumption to estimate the cohort variance of consumption. In this Section, we describe the data and the way we construct the three ingredients of our test: consumption inequality, income inequality and the income shocks.

The first step of our analysis is to construct the variance of log consumption and log income at the cohort level, our measures of consumption and income inequality, respectively. For this purpose, we use the SHIW, a representative sample of the Italian resident population conducted by the Bank of Italy.[18] The SHIW provides a measure of total non-durable consumption, not just food, thus overcoming one of the main limitations of other panels, such as the PSID, that have been used to test intertemporal consumption models. The survey also provides data on after-tax household disposable income, distinguishing between after-tax earnings, transfers and income from capital.[19]

From 1980 to 1984, the SHIW was conducted every year (with the exception of 1985), and every two years since 1987 (with the exception of a three-year interval between 1995 and 1998). Since 1986, it covered about 8,000 households, defined as groups of individuals related by blood, marriage or adoption and sharing the same dwelling. After 1987, SHIW has re-interviewed some households from the previous surveys. The panel component has increased over time: 15% of the sample was re-interviewed in 1989,

[17] There is little evidence on the plausibility of the classical measurement error assumption for consumption. Ahmed *et al.* (2006) compare diary and interview measures of food consumption in Canadian surveys (assuming that true consumption is the one that comes from the diary survey) and find evidence that the measurement error in food consumption is neither mean independent of true consumption nor homoscedastic.

[18] The survey is available online to all external users at http://www.bancaditalia.it. Questionnaire and documentation is available in English. Jappelli and Pistaferri (2010) report a detailed analysis of the quality of the SHIW data.

[19] Sampling is in two stages, first municipalities and then households. Municipalities are divided into 51 strata defined by 17 regions and three classes of population size (more than 40,000, 20,000–40,000, less than 20,000). Households are randomly selected from registry office records.

27% in 1991 and about 45% after 1993.[20] Response rates in the panel section of the SHIW are generally above 70%, in line with other microeconomic data sets.[21] Given the rotating sample structure, the number of repeated observations on households in our sample ranges from a minimum of two (households interviewed in two consecutive surveys), to a maximum of 10 (households interviewed each time from 1987 to 2006.

To minimise measurement error, we exclude cases in which the head changes over the sample period or gives inconsistent age figures. In most cases, the excluded households are those facing breaking-out events (widowhood, divorce, separation etc.), leading to changes in household head. Inconsistent age figures can reflect an unrecorded change in household head or measurement error. After these exclusions, the sample has about 50,000 consumption and income observations.

4.1. *The Variance of Income and Consumption*

We define consumption as the sum of all expenditure categories except durables. In our basic definition, we exclude services from durables but, in robustness checks, we experiment with a broader definition of consumption that includes housing services and one that is adjusted for equivalence scales. Income is defined as the sum of labour income and transfers of all household members, excluding income from capital (real and financial assets). These are the standard consumption and income concepts used in studies that test the implications of intertemporal consumption decisions.[22]

Figure 3 reports the variance of log consumption and log disposable income from 1980 to 2006.[23] All statistics are computed using sample weights. Jappelli and Pistaferri (2010) report that, by international standards, Italy has high income inequality, and that inequality is greater for earnings than for disposable income (net of non-financial income). They also report that demographic variables (age, family size, education, regional dummies) absorb about 40% of the income variability.

Over time, Figure 3 shows that there is a dramatic increase in income inequality, particularly during the 1991–93 recession (inequality increases by 50% between 1980 and 2006, after a 70% peak in 1998). The most plausible explanation for the increase in income inequality over the 1990s points to extensive labour market reforms, raising labour market instability. Indeed, during the decade, fixed term contracts were deregulated, widening their use, temporary work agencies were permitted and restrictions concerning fixed term contracts for unskilled workers were lifted. As a consequence, the overall index of Employment Protection Legislation constructed by the OECD declined from 3.6 in the late 1980s, to 2.7 to in the late 1990s and to 1.9 in

[20] In the panel component, the sampling procedure is also determined in two stages: selection of municipalities (among those sampled in the previous survey) and selection of households re-interviewed. This implies that there is a fixed component in the panel (for instance, households interviewed 10 times between 1987 and 2006, or four times from 2000 to 2006) and a new component every survey (for instance, households re-interviewed only in 2006).

[21] For instance, the net response rate in the US CEX is slightly above 80% for the Interview and Diary samples.

[22] In Section 7, we check the robustness of the results using alternative definitions of income.

[23] For the descriptive analysis, we can rely on earlier surveys. However, since the SHIW panel was first introduced in 1989, we cannot estimate the variance of income shocks in 1980–86, and so our regression analysis is limited to the 1987–2006 period.

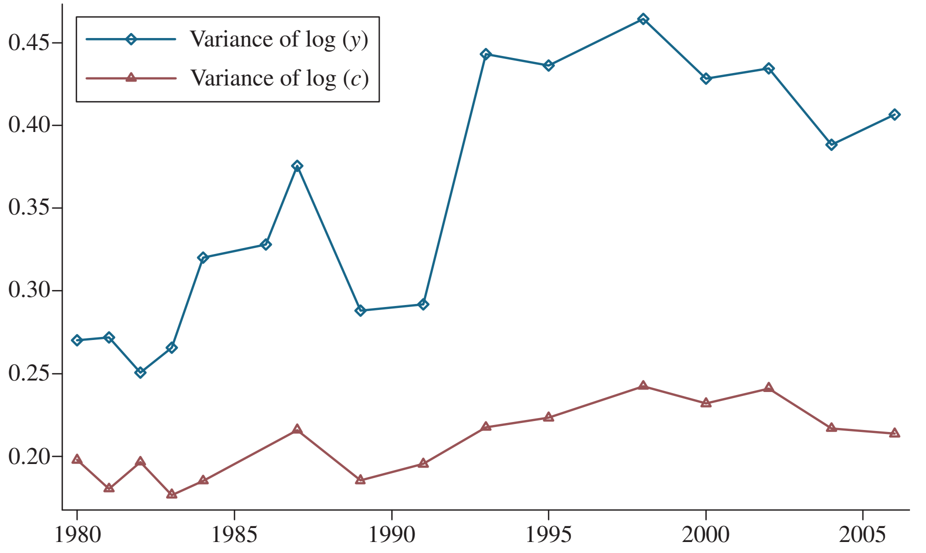

Fig. 3. *Income and Consumption Inequality,* 1980–2006

2003. According to this explanation, the increase in inequality is primarily of a transitory nature (raising income instability).

Figure 3 also shows that inequality is lower for consumption than for income. Jappelli and Pistaferri (2010) find that the level of consumption inequality is higher for the definition of non-durable consumption net of housing rents. Demographic variables also absorb part of the income variability. Over time, consumption inequality increases in the last decade but much less than income inequality: from 1980 to 2006 inequality increases by about 10%, after reaching a 20% peak in 2002.

Summing up, the increase in income inequality is not matched by a parallel increase in consumption inequality. Equation (8) shows that such divergence can be accounted for by changes in the insurance and excess sensitivity parameters over time (in particular, around the 1999 EMU policy shift), or by changes in the relative importance of transitory and permanent income shocks.

Our analysis is performed at the cohort level and, in our basic estimates, we exclude households headed by individuals older than 60 years or younger then 25 (regardless of year of birth). These exclusions are motivated by concern over two sources of potential sample bias. The first exclusion arises from the different earnings processes experienced by young and old households and from the different determinants of income shocks in old age. Furthermore, it is well known that survival probabilities tend to be positively correlated with income, especially in old age, inducing sample selection. The second source of potential bias is a correlation between income and young household heads peculiar to our sample. In Italy, young working adults with independent living arrangements tend to be wealthier than average, because most young working adults live with their parents.[24] Excluding individuals younger than 25 also

[24] For instance, the fraction of income recipients below 30 years of age is about 20%, whereas the fraction of household heads in that age bracket is less than 10%.

implies that we include only people who have completed school, an important consideration when we group households on the basis of education or use schooling to remove the permanent component of inequality.

We use the repeated cross-sections to sort the data by the year of birth of the head of the household. The first cohort includes all households whose head was born in 1930, the second those born in 1931 and so on up to the last cohort, including those born in 1970 (for robustness, we also present results with three and five-years cohorts). We remove the demographic component of inequality regressing log income and log consumption on age, education, gender, family size, number of kids, area of residence and year dummies.

Figure 4 displays var[ln($y_{i,a,t}$)] and var[ln($c_{i,a,t}$)] of six cohorts born between 1936 and 1965 (notice that the cohort born in 1961–5 is observed only after 1985). The results confirm the three stylised facts emerging from the aggregate evidence in Figure 3: cohort income inequality is substantially higher than consumption inequality; for each cohort, there is a dramatic increase in income inequality in the early 1990s (especially for the earlier cohorts) and a decline in the later part of the sample; there is also an increase in consumption inequality for most cohorts, but the dynamics of consumption inequality are much smoother than those of income.

4.2. *The Variance of Permanent and Transitory Income Shocks*

The next step of our analysis is to compute the time series of the variances of the permanent and transitory shocks for each cohort, using the panel section of the SHIW 1987–2006. As explained in Section 2, we obtain non-parametric estimates of var($e_{i,a,t}$) and var($u_{i,a,t}$) using (9) and (10), which assume that income is the sum of a random walk permanent component and a serially uncorrelated transitory component.

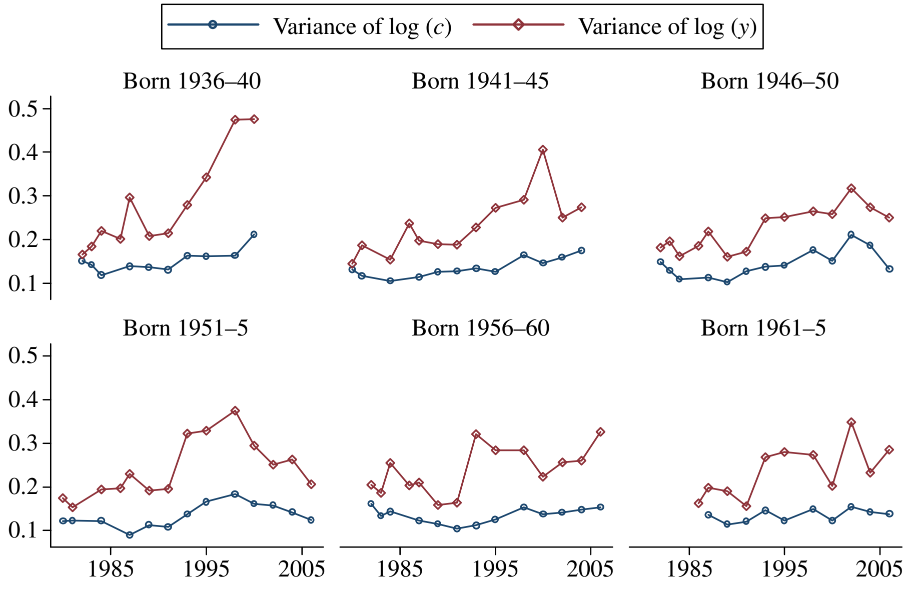

Fig. 4. *Consumption and Income Inequality by Selected Cohorts*

To check the consistency of the estimated income process with the model in equation (1), note that the income process implies the following testable restrictions on the autocovariance matrix of the first difference of income growth (neglecting for simplicity the contribution of demographics):

$$E(\Delta y_{i,a,t}\Delta y_{i,a-j,t-j}) = \begin{cases} \mathrm{var}(u_{i,a,t}) + \mathrm{var}(e_{i,a,t}) + \mathrm{var}(e_{i,a-1,t-1}) & \text{for } j = 0 \\ -\mathrm{var}(e_{i,a-1,t-1}) & \text{for } j = 1 \\ 0 & \text{for } j > 1 \end{cases}.$$

Pooling data for all years, we find that the estimated autocovariance at the first order ($j = 1$) is -0.0798 (SE 0.0050) but autocovariances after the first order drop abruptly to zero and are statistically insignificant.[25] Hence, we conclude the data are consistent with the specified income process and inconsistent with income processes with an AR component (where the decline after the first order is slower) or a random growth component (where autocovariances are positive and significant even at long lags).

Since the SHIW is conducted every two years (with a three-year gap in 1995–8), there is a slight complication in estimating the variance of the income shocks. We show in the Appendix that our estimates of the variance of permanent shocks are valid even in the presence of sample gaps; however, the estimates of the variance of transitory shocks are missing in the years in which the survey is not conducted. We solve the problem assuming that $\mathrm{var}(e_{i,a,t})$ can be approximated by a smooth function of adjacent variances, and reformulate accordingly the estimated (7) and (8).[26]

Over the sample period, we estimate that all cohorts experience an increase in the variance of transitory income shocks in the earlier part of the sample. This is an indication of increased income instability and, as we discussed in Section 4.1, it is likely to derive from the labour market reforms and the associated greater labour flexibility. At the same time, we estimate that there is no increase in the variance of permanent shocks (and even a decline, especially for the cohorts born in 1946–50 and 1951–5). The combined evidence suggests that the increase in income inequality that we observe in the early 1990s is mainly attributable to an increase in the transitory component of inequality.

5. Empirical Results

In this Section, we merge data on $\mathrm{var}[\ln(y_{i,a,t})]$, $\mathrm{var}[\ln(c_{i,a,t})]$ obtained from cross-sectional data with data on $\mathrm{var}(e_{i,a,t})$ and $\mathrm{var}(u_{i,a,t})$ obtained from the panel and report estimates of the parameters of (8). In Table 2, we report estimates of the parameters of (8) reformulated as (A.6) in the Appendix. The estimates of the parameters are obtained by non-linear least squares correcting the standard errors for heteroscedasticity of unknown form.[27] Since each cohort is defined over one-year cells, the sample size includes 185 observations, corresponding to a maximum of nine observations for each cohort;

[25] For example, the second-order autocovariance is -0.0044 (SE 0.0040) and the third-order autocovariance is -0.0016 (0.0045).

[26] We take into account the two-year gap issue when estimating the income autocovariances discussed above.

[27] We adopt this more general heteroscedasticity correction, which does not require the specification of a particular form of heteroscedasticity (such as one that depends on cohort cell size). We also report estimates of ϕ and ψ, instead of the 'reduced form parameters' because they have a more natural metric – for example, ϕ is ideally bounded between 0 (the full insurance case) and 1 (the PIH case).

Table 2

Difference-in-Difference of var(y) *and* var(c), *One-Year Cohorts*

	Total sample			Excluding $n < 30$		
ϕ	0.989	1.089	0.987	0.861	0.947	0.847
	(0.056)**	(0.061)**	(0.057)**	(0.114)**	(0.136)**	(0.117)**
ψ	0.282	0.244	0.295	0.224	0.053	0.114
	(0.081)**	(0.116)*	(0.097)**	(0.139)	(0.691)	(0.320)
ϕ_E		−0.378			−0.271	
		(0.152)*			(0.266)	
ψ_E		0.030	−0.037		0.374	0.268
		(0.178)	(0.175)		(0.705)	(0.348)
Observations	185	185	185	82	82	82

Notes. The table reports non-linear least squares estimates of various versions of equation (8) in the text. 'Excluding $n < 30$' restricts the sample to cells with at least 30 observations when computing the variances of income shocks in the 1987–2006 panel. Robust standard errors are reported in parenthesis. * Significance at 1% level; ** significance at the 5% level.

cells where the income shocks are computed on less than five observations are dropped. In the baseline specification, we do not distinguish between pre and post-EMU observations. The results reported in the first column show that the insurance parameter is $\phi = 0.989$ and the excess sensitivity parameter is $\psi = 0.282$. Thus, in our total sample estimate, we find evidence that consumers do not smooth permanent shocks and that consumption reacts also to transitory shocks although much less than one-for-one.[28]

The evidence for excess sensitivity is broadly consistent with previous studies on the effect of transitory income shocks on consumption expenditure. Using CEX quarterly panel data, Souleles (1999) and Parker (1999) examine, respectively, the response of household consumption to income tax refunds and to predictable changes in Social Security with holdings. Souleles finds evidence that the marginal propensity to consume is at least 35% of refunds within a quarter and Parker finds that consumption reacts significantly to changes in tax rates. Jappelli and Pistaferri (2006), using data from the 1989–95 SHIW, estimate the parameters that minimise the distance between the empirical and the theoretical transition matrix of the consumption distribution and also find evidence that the response of consumption to transitory shocks is larger than predicted by the PIH. Browning and Crossley (2001) survey several other studies reporting evidence that consumption overreacts to anticipated income innovations.

In the second column of Table 2, we let the insurance and excess sensitivity parameters vary over time. The hypothesis that we test is that consumption has become less sensitive to income shocks after the introduction of the euro (a negative value of ϕ_E and ψ_E). The estimates indicate that ϕ_E is negative (−0.378) and statistically different from zero at the 5% level after the introduction of the euro, whereas ψ_E is close to zero and insignificant. A formal statistical test of the joint hypothesis of no EMU effect rejects the null hypothesis $\phi_E = \psi_E = 0$ at the 5% confidence level. In the third specification of Table 2, we constrain the insurance coefficient to be constant over time

[28] Recall from footnote 15 that in the presence of measurement error in income, parameter ψ is upward biased. The estimate in the first column of Table 2 would be consistent with full insurance of transitory shocks ($\psi = 0$) if $\lambda = 0.92$; that is, if measurement error in income accounted for 8% of the total variance of the mean-reverting income component.

($\phi_E = 0$). The excess sensitivity coefficients are hardly affected, showing only a slight reduction after the introduction of the euro.

The last three columns repeat the estimation including cells where the income shock variances are more reliable because they are computed from cells with at least 30 observations, reducing the sample size to 82 observations. The results only partly confirm the findings obtained for the total sample. The insurance coefficient is statistically different from zero at the 1% level ($\phi = 0.947$), whereas the other coefficients are imprecisely estimated. There is no evidence that the excess sensitivity coefficient declines in the post-1999 period. If anything, the point estimate of ψ_E is positive. In this restricted sample, therefore, a formal test does not reject the hypothesis of no EMU effect.

In Table 3, we define cohorts on the basis of three years of birth, expanding considerably the number of observations on which we compute the income shocks but reducing the number of cells. The results are qualitatively unaffected. In the total sample, we find again a reduction in the insurance parameter and a slight increase in the excess sensitivity parameter after the introduction of the euro. Restricting the sample to observations drawn from cells with at least 30 households shows that these effects are statistically insignificant, so that the hypothesis of no EMU effect is not rejected at conventional statistical levels.

A further experiment we perform is to split the sample between households in which the head has completed college and those in which he or she has not, and then compute the variance of income shocks on the basis of cohorts defined over five-year intervals.[29] This reduces the number of valid observations that we use to estimate the two parameters of (8) but allows estimation of different income processes for households with different levels of education.

Figure 5 reports income and consumption inequality from 1980 to 2006 for the two groups of households and three selected cohorts, and shows that there is a much

Table 3

Difference-in-Difference of var(y) *and* var(c), *Three-Year Cohorts*

	Total sample			Excluding $n < 30$		
ϕ	1.103 (0.035)**	1.130 (0.034)**	1.102 (0.036)**	0.801 (0.118)**	0.869 (0.139)**	0.802 (0.119)**
ψ	0.205 (0.133)	0.184 (0.168)	0.218 (0.148)	0.273 (0.095)**	0.231 (0.140)	0.256 (0.122)*
ϕ_E		−0.481 (0.216)*			−0.199 (0.273)	
ψ_E		0.121 (0.232)	−0.051 (0.330)		0.107 (0.196)	0.050 (0.181)
Observations	87	87	87	53	53	53

Notes. The Table reports non-linear least squares estimates of various versions of (8) in the text. 'Excluding $n < 30$' restricts the sample to cells with at least 30 observations when computing the variances of income shocks in the 1987–2006 panel. Robust standard errors are reported in parenthesis. * significance at 1% level; ** significance at the 5% level.

[29] Since in Italy, the number of college graduates is, on average, only 10% per cohort, we cannot define cohorts on the basis of college education. The sample with high-school and college accounts for 60% of the total.

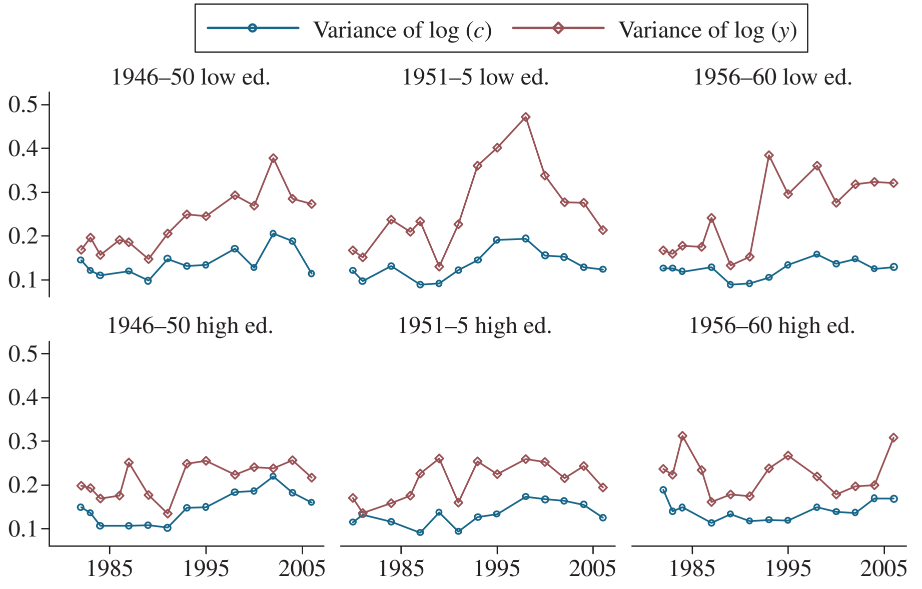

Fig. 5. *Consumption and Income Inequality by Selected Cohorts and Education Groups*

stronger increase in inequality among households with relatively low education. The estimates of separate income processes for the two groups show that the increase of income inequality is largely accounted for by an increase in the variance of transitory shocks among households who have not completed college. This is further evidence that the increase in income inequality in the last decade is because of the liberalisation of the labour market and the associated increase in temporary and part-time contracts, which are more prevalent among blue-collar workers.

The regression results reported in Table 4 show some differences between the two groups regarding the parameters ϕ and ψ. The insurance parameter is higher among households with lower education (0.893) than among those who completed college (0.720), suggesting that people with higher education can more easily smooth permanent income fluctuations. When we distinguish between pre and post-EMU samples, we find that in both groups there is a drop in the excess sensitivity parameter after the introduction of the euro but the statistical test never supports the hypothesis that the EMU has increased consumption smoothing.

6. Robustness Analysis

In this Section, we check if our results are robust to different definitions of income, alternative measures of financial integration and restrictions that the theory imposes on the joint behaviour of consumption and income. We also compare statistically the dynamics of consumption and income inequality in Italy with the dynamics of the same variables in the UK. All the regressions in this Section use cells defined on the basis of one-year cohorts.

Table 4

Difference-in-Difference of var(y) *and* var(c), *Five-Year Cohorts by Education*

	No college education			College education		
ϕ	0.893	0.900	0.910	0.720	0.788	0.673
	(0.125)**	(0.171)**	(0.127)**	(0.151)**	(0.219)**	(0.168)**
ψ	0.327	0.356	0.354	0.337	0.421	0.451
	(0.109)**	(0.115)**	(0.110)**	(0.090)**	(0.137)**	(0.119)**
ϕ_E		0.023			−0.221	
		(0.261)			(0.348)	
ψ_E		−0.209	−0.187		−0.160	−0.186
		(0.568)	(0.442)		(0.198)	(0.185)
Observations	32	32	32	31	31	31

Notes. The Table reports non-linear least squares estimates of various versions of (8) in the text. The sample is restricted to cells with at least 30 observations when computing the variances of income shocks in the 1987–2006 panel. Robust standard errors are reported in parenthesis. Robust standard errors are reported in parenthesis. * significance at 1% level; ** significance at the 5% level.

6.1. *Alternative Measure of Financial Integration*

Recall that in (8), we have assumed that financial integration affects the dynamics of income and consumption inequality only after 1999. The assumption is questionable, as the path to financial integration started well before 1999, as highlighted by the dynamics of the Italian-German interest rate spread in Figure 1. Accordingly, we use the spread itself (S) as a specific measure of financial integration:

$$\begin{aligned}\Delta\text{var}(\ln y_{i,a,t}) - \Delta\text{var}(\ln c_{i,a,t}) &= [1 - (\phi + \phi_E S)^2]\text{var}(u_{i,a,t}) \\ &\quad + [1 - (\psi + \psi_E S)^2]\text{var}(e_{i,a,t}) - \text{var}(e_{i,a,t-1}). \qquad (11)\end{aligned}$$

Note that since a reduction in the spread signals greater financial integration, if the EMU has reduced the sensitivity of consumption to income shocks, we should expect in this specification $\phi_E > 0$ and $\psi_E > 0$. Comparison of (8) and (11) is useful because it allows us to isolate the effect of changes in credit and insurance market opportunities after the introduction of the euro.

The results are reported in the first two columns of Table 5 for the total sample and for the sample excluding cells with less than 30 observations, respectively. In both regressions, the point estimates of ϕ_E and ψ_E are positive but not statistically different from zero, confirming the results in Section 5.

In a related test, we split the sample between regions that are more and less financially backward according to the indicator developed in Guiso *et al.* (2004*b*). The indicator is based on a set of questions contained in the SHIW on whether households were denied credit or discouraged from borrowing. The results, not reported for brevity, do not highlight differential responses of consumption after the introduction of the EMU in regions with different degrees of financial development.

6.2. *Alternative Income Measures*

It may be argued that our measure of income net of taxes and transfers already incorporates some smoothing, as provided by the tax system, government insurance or

Table 5

Robustness Analysis

	Using the spread as a measure of integration		Using family earnings		Using covariance restrictions		Using UK as a comparison	
	Total sample	Excluding $n < 30$	Total sample	Excluding $n < 30$	Total sample	Excluding $n < 30$	Total sample	Excluding $n < 30$
ϕ	0.639	0.921	0.899	0.666	0.861	0.593	0.955	0.922
	(0.176)**	(0.284)**	(0.075)**	(0.248)**	(0.054)**	(0.130)**	(0.072)**	(0.112)**
ψ	0.033	−0.009	−0.130	0.045	−0.475	−0.177	−0.088	0.319
	(0.016)*	(0.041)	(0.153)	(0.311)	(0.144)**	(0.220)	(0.204)	(0.279)
ϕ_E	0.028	0.006	0.263	0.339	−0.135	0.119	0.302	0.210
	(0.282)	(0.412)	(0.120)*	(0.150)*	(0.072)	(0.114)	(0.085)**	(0.177)
ψ_E	0.027	0.029	0.250	0.165	0.351	0.279	0.006	0.300
	(0.025)	(0.041)	(0.146)	(0.171)	(0.110)**	(0.144)	(0.236)	(0.338)
$\phi_{ITA,E}$							−0.164	−0.581
							(0.224)	(0.321)
$\psi_{ITA,E}$							−0.027	−0.070
							(0.246)	(0.309)
Observations	185	82	180	61	185	82	282	160

Notes. The Table reports non-linear least squares estimates of the sensitivity parameters. We define cells on the basis of one-year cohorts. In columns 1 and 2, we use the Italy-Germany interest rate spread as a measure of financial integration, and estimate (11) in the text. In columns 3 and 4, we use family earnings as income measure. In columns 5 and 6, we use (8) and (12) jointly to estimate the consumption parameters. In the last two columns, we estimate (13) in the text, using observations from both Italy and the UK. 'Excluding $n < 30$' restricts the sample to cells with at least 30 observations when computing the variances of income shocks in the Italian and British panels. Robust standard errors are reported in parenthesis.
* Significance at 1% level; ** significance at the 5% level.

private transfers. To check whether our results are sensitive to this criticism, in the third and fourth columns of Table 5, we estimate the income process and compute the income shock variances by using a definition of income that excludes public and private transfers. We again obtain sharp estimates of the insurance parameters but no appreciable change in the pattern of the parameter estimates after the introduction of the euro. If anything, we find that using this alternative income measure, the sensitivity of consumption to permanent shocks has increased (rather then decreased) after the introduction of the euro.

6.3. *Using Additional Covariance Restrictions*

When income and consumption are available in the same dataset, as in the Italian case, one can also use the restrictions implied by the theory on the covariance between income and consumption. In fact, note that the within cohort covariance of log income (1) and log consumption (3) is:

$$\mathrm{cov}(\ln c_{i,a,t}, \ln y_{i,a,t}) = \mathrm{cov}(\ln c_{i,a_0,t-a+a_0}, \ln \pi_{i,a_0,t-a+a_0}) + \phi \sum_{j=a_0+1}^{a} \mathrm{var}(u_{i,j,t-a+j}) + \psi\, \mathrm{var}(e_{i,a,t}).$$

Taking first differences of the cohort-specific covariance yields an expression that depends on the same parameters and regressors as (8):

$$\Delta \mathrm{cov}(\ln c_{i,a,t}, \ln y_{i,a,t}) = \phi \mathrm{var}(u_{i,a,t}) + \psi \Delta\, \mathrm{var}(e_{i,a,t}). \tag{12}$$

We therefore estimate (8) and (12) jointly imposing cross-equation restrictions and report the results in the fifth and sixth columns of Table 5. Comparing the results with columns 2 and 5 of Table 2, the estimates point to slightly lower sensitivity of consumption to permanent shocks. In both specifications, however, we find no evidence that ϕ and ψ have statistically significantly decreased after the introduction of the euro.

6.4. *Using UK Cohorts as a Control Group*

Our identification strategy relies on the assumption that the only reason why ϕ and ψ change after 1999 is the introduction of the euro. This assumption is questionable, as the process of financial integration has been a global phenomenon not just confined to the euro-area, potentially leading to spurious results (in particular, the stability of ϕ and ψ over time might result from offsetting effects). To account for this possibility we expand the analysis considering the UK as a control country experiencing the global and EU-related integration process but not the specific euro-effect. Indeed, Figure 1 shows that interest rate convergence has been significantly stronger in Italy than in the UK. In terms of interest rate convergence, we therefore take these two countries as representative of the group of euro adopters and non-adopters.

We use UK microeconomic data from the 1991–2004 FES and 1991–2004 BHPS, select a sample with similar characteristics (in particular, we restrict the age of the household head in the 25–60 interval) and regress log income and consumption on the same characteristics used in the Italian case. We then form cohort-level variances of log

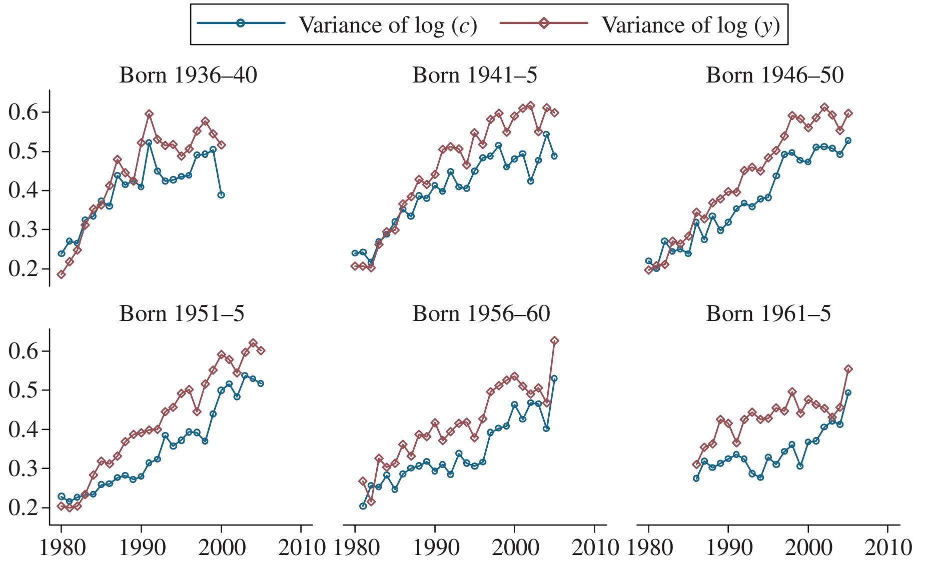

Fig. 6. *Consumption and Income Inequality by Selected Cohorts in the UK*

income and consumption (defining cohorts on the basis of one-year-cells) and estimate the variances of transitory and permanent income shocks using the same income process.[30] Figure 6 is the equivalent of Figure 4 for the UK. It plots variances of log consumption and log income for six selected cohorts. As in the Italian case, the Figure shows a growing detachment between income and consumption inequality for all cohorts considered and particularly for the younger cohorts.

We then pool the Italian and UK data and estimate an extended version of (8) which allows for post-1999 effect common to both countries (ϕ_E and ψ_E) and a post-1999 effect specific only to Italy ($\phi_{ITA,E}$ and $\psi_{ITA,E}$):[31]

$$\begin{aligned}\Delta\mathrm{var}(\ln y_{i,a,t}) - \Delta\mathrm{var}(\ln c_{i,a,t}) &= [1 - (\phi + \phi_E E + \phi_{ITA,E}\mathrm{ITALY} \times E)^2]\mathrm{var}(u_{i,a,t}) \\ &\quad + [1 - (\psi + \psi_E E + \psi_{ITA,E}\mathrm{ITALY} \times E)^2]\mathrm{var}(e_{i,a,t}) - \mathrm{var}(e_{i,a,t-1}). \qquad (13)\end{aligned}$$

The results are reported in the last two columns of Table 5. In accordance with previous estimates, there is no evidence that the sensitivities to permanent or transitory shocks have declined in Italy more than in the UK after 1999 (if anything, the sensitivity with respect to permanent shocks increases in the later part of the sample in both countries).

[30] A full description of the two UK datasets and sample selection is contained in Blundell and Etheridge (2010).

[31] We experiment also with a less parsimonious specification, including the terms $\phi_{ITA} \times$ ITA and $\psi_{ITA} \times$ ITA, capturing the possibility that smoothing and insurance opportunities are country-specific even in the absence of a process of financial integration. We do not reject the hypothesis that the two additional coefficients are statistically different from zero.

6.5. *Additional Tests*

To further check the robustness of the results, we perform a number of additional sensitivity checks: since the process of European financial integration has preceded the introduction of the euro, we test the stability of the parameters defining the EMU sample as 1996–2006 or 1998–2006; in contrast, since financial integration might impact consumption with a lag, we test the stability of the results by restricting the EMU sample to the years 2002–6; we define cohorts on the basis of cells defined over 7 or 10 years, to check that aggregating over cells does not bias our results; we subtract from consumption imputed rents that are likely to be subject to substantial measurement error; we deflate consumption and income by the OECD equivalence scale.[32] These experiments confirm the patterns found in Tables 2–4. The point estimates of the insurance and excess sensitivity parameters generally decline after the introduction of the euro but the hypothesis of no EMU effect is not rejected at standard confidence levels.

7. Conclusions

In this article, we present a new empirical strategy for testing if financial integration improves risk-sharing opportunities and consumption smoothing. Our test is based on a decomposition of the variance of consumption growth into a component that depends on the variance of permanent income shocks and one that depends on the variance of transitory shocks. We then test if the process of financial market integration and liberalisation brought about by the introduction of the euro has made consumption less sensitive to income shocks in Italy. The article makes a significant contribution also from a methodological point of view. We use panel data on income to identify non-parametrically a time series of the variances of the income shocks. We then rely on repeated cross-sections of consumption and income to identify the degree of smoothing with respect to income shocks and test if it has declined after the introduction of the euro.

In the data, we uncover a divergence between consumption and income inequalities: in particular, that the dramatic increase in income inequality has not been matched by an increase in consumption inequality. Our point estimates of the effect of permanent and transitory shocks support the PIH (an insurance parameter close to one), although in most specifications, we find that also transitory shocks impact consumption (an excess sensitivity parameter in the order of 0.2–0.3). We also find that the point estimates of the insurance and excess sensitivity parameters tend to decline after the introduction of the euro but statistically the null hypothesis of no EMU effect is not rejected at standard confidence levels. The result is robust to the presence of classical measurement error in income and consumption, alternative measures of financial integration and different measures of income and consumption. We also compare the dynamics of income and consumption inequality using comparable microeconomic data for the UK and exploit the additional restrictions imposed by the theory on the

[32] The OECD equivalence scale is defined as: $E = 1 + 0.5 \times$ (number of children) + $0.7 \times$ (number of adult members $- 1$). A child is any household member aged 16 or less.

covariance between income and consumption. All of these robustness checks confirm the baseline results.

We conclude that during our sample period, the ability of consumers to smooth income shocks has not changed and that the diverging trends between income and consumption inequality is explained by the fact that the increase in income inequality is primarily because of an increase in transitory inequality. Since consumers smooth transitory shocks to a much larger extent than permanent shocks, the increase in income inequality has not translated one-for-one into an increase in consumption inequality.

The lack of decline of excess sensitivity of consumption after the introduction of the euro signals that financial integration in Europe is a slow process, which so far has not produced significant changes in consumption smoothing opportunities. As highlighted by the European Commission (2008), financial integration remains a work in progress for the euro area. While integration has progressed substantially since, and in part owing to, the introduction of the euro, many markets are still fragmented and the pace of integration varies among Member States. Indeed, the effect of financial market integration is quite visible in the European bond markets, and there is some evidence of increased integration of equity markets, with a decline in home bias, although important institutional barriers remain. Credit markets, by contrast, have integrated at a slower pace, reflecting in part the informational advantage enjoyed by local lenders, and differences in regulation, taxes and labour regulation. We speculate that further progress towards credit market integration is necessary to feel the benefits of integration and its effect on consumption smoothing.

Appendix

In this Appendix, we discuss how we deal with the fact that the survey is conducted every other year, and that there is a three-year gap between the 1995 and 1998 surveys. Our starting points are equations (2) and (4) in the main text, which we re-propose here:

$$\mathrm{var}(\ln y_{i,a,t}) = \mathrm{var}(\pi_{i,a_0,t-a+a_0}) + \sum_{j=a_0+1}^{a} \mathrm{var}(u_{i,j,t-a+j}) + \mathrm{var}(e_{i,a,t}), \tag{A.1}$$

$$\mathrm{var}(\ln c_{i,a,t}) = \mathrm{var}(\ln c_{i,a_0,t-a+a_0}) + \phi^2 \sum_{j=a_0+1}^{a} \mathrm{var}(u_{i,j,t-a+j}) + \psi^2 \sum_{j=a_0+1}^{a} \mathrm{var}(e_{i,j,t-a+j}). \tag{A.2}$$

The expressions for time $t-2$ (age $a-2$) are:

$$\mathrm{var}(\ln y_{i,a-2,t-2}) = \mathrm{var}(\pi_{i,a_0,t-a+a_0}) + \sum_{j=a_0+1}^{a-2} \mathrm{var}(u_{i,j,t-a+j}) + \mathrm{var}(e_{i,a-2,t-2}), \tag{A.3}$$

$$\mathrm{var}(\ln c_{i,a-2,t-2}) = \mathrm{var}(\ln c_{i,a_0,t-a+a_0}) + \phi^2 \sum_{j=a_0+1}^{a-2} \mathrm{var}(u_{i,j,t-a+j}) + \psi^2 \sum_{j=a_0+1}^{a-2} \mathrm{var}(e_{i,j,t-a+j}). \tag{A.4}$$

The differences between equations (A.1) and (A.3) and between equations (A.2) and (A.4) are, respectively:

$$\Delta^2\mathrm{var}(\ln y_{i,a,t}) = \mathrm{var}(u_{i,a,t}) + \mathrm{var}(u_{i,a-1,t-1}) + \Delta^2\mathrm{var}(e_{i,a,t}),$$
$$\Delta^2\mathrm{var}(\ln c_{i,a,t}) = \phi^2[\mathrm{var}(u_{i,a,t}) + \mathrm{var}(u_{i,a-1,t-1})] + \psi^2[\mathrm{var}(e_{i,a,t}) + \mathrm{var}(e_{i,a-1,t-1})].$$

Finally, the equivalence of the difference-in-difference expression (7) that we use as a basis for estimation is:

$$\begin{aligned}\Delta^2\mathrm{var}(\ln y_{i,a,t}) - \Delta^2\mathrm{var}(\ln c_{i,a,t}) = {} & (1-\phi^2)[\mathrm{var}(u_{i,a,t}) + \mathrm{var}(u_{i,a-1,t-1})] + (1-\psi^2)\mathrm{var}(e_{i,a,t}) \\ & - \psi^2\mathrm{var}(e_{i,a-1,t-1}) - \mathrm{var}(e_{i,a-2,t-2}).\end{aligned} \tag{A.5}$$

Using extensions of equations (9) and (10), in panel data we can identify non-parametrically:

$$\mathrm{E}[(\ln y_{i,a,t} - \ln y_{i,a-2,t-2})(\ln y_{i,a+2,t+2} - \ln y_{i,a-4,t-4})] = \mathrm{var}(u_{i,a,t}) + \mathrm{var}(u_{i,a-1,t-1}),$$

$$-\mathrm{E}[(\ln y_{i,a,t} - \ln y_{i,a-2,t-2})(\ln y_{i,a+2,t+2} - \ln y_{i,a,t})] = \mathrm{var}(e_{i,a,t}),$$

$$-\mathrm{E}[(\ln y_{i,a-2,t-2} - \ln y_{i,a-4,t-4})(\ln y_{i,a,t} - \ln y_{i,a-2,t-2})] = \mathrm{var}(e_{i,a-2,t-2}).$$

However, $\mathrm{var}(e_{i,a-1,t-1})$ remains not identified. We assume that it can be approximated by a smooth function of adjacent variances. In this specific case, we assume:

$$\mathrm{var}(e_{i,a-1,t-1}) = \frac{\mathrm{var}(e_{i,a,t}) + \mathrm{var}(e_{i,a-2,t-2})}{2}$$

and hence rewrite (A.5) as:

$$\begin{aligned}\Delta^2\mathrm{var}(\ln y_{i,a,t}) - \Delta^2\mathrm{var}(\ln c_{i,a,t}) = {} & (1-\phi^2)[\mathrm{var}(u_{i,a,t}) + \mathrm{var}(u_{i,a-1,t-1})] \\ & + \left(1-\psi^2-\frac{\psi^2}{2}\right)\mathrm{var}(e_{i,a,t}) - \left(1+\frac{\psi^2}{2}\right)\mathrm{var}(e_{i,a-2,t-2}).\end{aligned} \tag{A.6}$$

This is the basic regression we run and whose results are reported in the first and fourth columns of Tables 2–4. We use a similar strategy to deal with the three-year gap between the 1995 and 1998 surveys.

Università di Napoli Federico II, CSEF and CEPR
Stanford University, NBER and CEPR

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