Financial Integration and Consumption Smoothing

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Abstract

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 liberalization brought about by the introduction of the euro has made consumption less sensitive to income shocks in Italy. The paper 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. Our procedure does not require that consumption and income are available in the same panel data. It can therefore be applied in all countries in which repeated cross-sectional consumption data can be combined with panel data on income.

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1. Introduction

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 to that of the United States. To what extent has this process of regulatory reform affected the ability of households to diversify, insure, and shoulder risks? This paper 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 liberalization. Applying Cochrane (1991) and Mace (1991) seminal contributions to aggregate data, Sorensen, Wu, Yosha and Zhu (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 due to individual-specific shocks, rather than to region or countrywide shocks.

In this paper we develop a new empirical strategy for testing if financial liberalization improves consumers' ability to wedge against income shocks. Our analysis is performed at the cohort level. We build on work by Deaton and Paxson (1995), Blundell and Preston (1998) and Blundell, Pistaferri and Preston (2008), and decompose the change in the variance of consumption into a component that depends on the variance of permanent 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 liberalization brought about by the introduction of the euro has affected the sensitivity of consumption to income shocks. The test allows us to recover two structural 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. We extend the frontier of current literature also considering explicitly the effect of removing borrowing constraints for the dynamics of the variance.

The paper 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. Instead, we rely on repeated cross-sectional data to construct the 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. It 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 United States and the United Kingdom provide compelling examples of such situation.

The paper is organized as follows. Section 2 reviews the literature on the effect of financial market integration on risk sharing opportunities and consumption smoothing. Section 3 discusses the macroeconomic developments in the euro-zone and Italy before and after the introduction of the euro. Section 4 explains how changes in the variance of consumption over time can signal changes in consumption smoothing. Section 5 presents the data, which are drawn from the 1987-2006 Survey of Household Income and Wealth, and explains how we construct the three ingredients of our test: consumption inequality, income inequality, and the variance of the income shocks. Section 6 presents the main empirical results, and checks their robustness using various definitions of consumption. Section 7 concludes.

2. Financial market integration, risk sharing and consumption smoothing

Economic theory predicts that the process of financial market integration should facilitate 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 recognized and applied to United States microeconomic data by Cochrane (1991) and Mace (1991), and later brought to bear on macroeconomic data by Obstfeld (1994), van 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 United States data for 1963-90, Asdrubali, Sorensen and Yosha (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, Sorensen and Yosha (1996) find that in the United States 39 percent of the shocks are absorbed via capital market smoothing, 13 percent via the fiscal channel and 23 percent via the credit market, while the remaining 25 percent are not smoothed. Sorensen and Yosha (1998) and Kalemli-Ozcan, Sorensen and Yosha (2004) 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

[1] See Jappelli and Pagano (2008) for a survey of the real effects of financial market integration in the context of the EMU.

percent, is much larger than in the United States. 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. Rubini, Parisi-Capone and Menegatti (2007) extend the analysis to 2006, and find that risk sharing in the EMU is still significantly lower than in the United States, but that it has significantly improved over time in the euro-zone and during the EMU period.[2]

These results must be taken with considerable caution. When applied to aggregate data, risk-sharing tests require highly unrealistic assumptions, which neglect the heterogeneity of the population within each country. In particular, the implicit assumption is that there is a representative agent within the economy (or region of a country), which is tantamount to assume that agents are fully insured against person-specific shocks (such as unemployment, low productivity due to health shocks, disability, etc.). However, the hypothesis of full insurance of idiosyncratic shocks is typically rejected using micro data, see Attanasio and Davis (1996) and Jappelli and Pistaferri (2006). Even introspectively, it is difficult to believe that insurance against idiosyncratic shocks, which are often private information, is easier to come by than insurance against country-specific shocks, which are mostly fully observable and cannot be manipulated.

In this paper we fill a gap in the literature and provide a test for the effect of financial integration on risk sharing opportunities based on microeconomic data. We assess with household level data how 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, 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 analyzed the impact of financial integration on consumption using household level data, empirical evidence with firm-level data exists.[3] Alfaro and Charlton

[2] Sørensen, Wu, Yosha and Zhu (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.

[3] Several studies using firm-level data document that financial development has a positive effect on access to finance and entry of new firms, see Guiso, Jappelli, Padula and Pagano (2004), Aghion, Fally and Scarpetta (2007) and Klapper, Laeven and Rajan (2005).

(2007) show that reducing restrictions on international capital flows enhances firm entry and other measures of entrepreneurship. Bertrand, Schoar and Thesmar (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, while firms in more bank-dependent sectors became more likely to undertake restructuring activities. Recent microeconomic evidence also throws light on the role that international financial integration can play in improving the allocation of capital across firms. Galindo, Schiantarelli and Weiss (2007) use firm-level panel-data from twelve Latin American countries to investigate whether capital account liberalization has increased the share of investment going to firms with a higher marginal return to capital. They develop an indicator of the efficiency of allocation of investment, and find that in the majority of cases financial liberalization has led to an increase in the efficiency with which investment funds are allocated. Since the sample spans pre and post-reform periods, this is one of the few studies that are able to show that this improvement in the allocation of domestic capital is actually traceable to the financial development occurring in the wake of the liberalization.

3. 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 homogenized. 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 United States. Jappelli and Pagano (2008) describe these developments, and the effect of the EMU on financial market integration, investment, growth, ability to response to macroeconomic shocks and risk sharing opportunities.

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, while other interest rates have converged rapidly. The convergence of the spreads toward zero is dramatic, particularly in the Italian case. Considering all EMU initial participants (and thus excluding Greece), the mean yield spread over the German yield fell from 218 basis points in 1995 to 20 in 1999. For Italy the convergence is even more dramatic, as the spread fell from 546 points in 1995 to 30 points in 1999.[4] The convergence is illustrated in Figure 1 with reference to the 10-year benchmark bonds (but qualitatively similar pictures are obtained for other maturities). The figure shows that the end-of-month cross-sectional standard deviation of the euro-area benchmark government bonds yields in the EMU countries relative to the 10-year German Bund fell from about 400 basis points in the early 1990s to less than 100 points after the introduction of the euro.

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 rised significantly between 1999 and 2007, reaching almost 30 percent (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.

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

[4] Most of the action derives from the convergence of the non-core EMU participants: Finland, Ireland, Italy, Portugal and Spain, and later Greece, which joined the euro area at the beginning of 2001. The bonds issued by Austria, Belgium, France and the Netherlands already featured low spreads relative to German bonds since 1996. This is because before EMU the probability of depreciation relative to the D-Mark was considerable in the first set of countries, but not in the second. Indeed, for the non-core EU countries the drop of the 10-year yield spreads is overwhelmingly due to the elimination of this possibility.

mortgage foreclosure, the length of trials, and judicial inefficiency in Italy are higher than in countries at a similar level of financial development.[5] Casolaro, Gambacorta and Guiso (2005) also stress that, compared to other countries, Italy features a lower level of social capital and trust, effecting 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-GDP ratio more than tripled from 9 percent in 1986 to almost 30 percent in 2006, and a 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.[6] The development of 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 industrialized 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.

[5] 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.

[6] Before 1986 the maximum statutory LTV ratio was 50 percent. In 1986 the LTV ratio for first-time-buyers was raised to 75 percent; the limit for repeat buyers was still 50 percent. In 1995 the maximum LTV was further raised to 80 percent.

4. The empirical strategy

We rely on the covariance restriction implied by the permanent income hypothesis to check if the variance of consumption is less closely correlated with the variance of income after the introduction of the euro. For this purpose, we rely on standard assumptions about the evolution of income shocks (see Deaton, 1991; Carroll, 1997; and Blundell and Preston, 1998). In particular, we decompose income into three parts: a deterministic component, a permanent component and a transitory shock:

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

where:

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

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 labor supply, bonuses, lottery prizes, and bequests. On the other hand, some of the innovations to earnings are highly persistent (non-mean reverting) and their effect cumulates over time. Examples of permanent innovations are generally associated with job mobility, promotions, lay-off, and severe health shocks. From previous work, we posit that the variance of permanent income shocks is 0.02, and that the variance of transitory income shocks is 0.04.

We assume that preferences between consumption and leisure are separable, and that the only relevant source of idiosyncratic uncertainty faced by consumers is after-tax household disposable income. Since insurance provided through government taxes and transfers and changing labor supply of the family passes through disposable income, in this paper we concentrate on the effect of financial integration on consumption smoothing rather than on income smoothing.

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

$$\ln y_{i,a,t} = x_{i,a,t}\beta + \pi_{i,a_0,t-a+a_0} + \sum_{j=a_0+1}^{a} u_{i,j,t-a+j} + e_{i,a,t}$$

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 labor 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:

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

where for notational convenience from now on we omit the subscript b for the variance terms. 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 don't need to assume that u and e are covariance-stationary.

Following Blundell, Pistaferri and Preston (2008), we obtain a similar decomposition of the variance of consumption, starting from an approximation of the Euler equation. To derive such approximation, let's 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. According to the permanent income model, consumption responds fully to permanent income shocks ($\phi\approx1$), while transitory shocks have negligible effects ($\psi\approx0$) because consumers use assets to smooth temporary income fluctuations. The buffer stock model delivers similar implications.[7] If there are complete markets, individual consumption is completely insulated from transitory as well as permanent shocks ($\phi=\psi=0$). Finally, models with partial insurance predict that consumers are able to insure also permanent shocks to a larger extent than in the PIH ($\psi\approx0$ and $0<\phi<1$).

The parameter ψ in equation (3) represents the extent to which consumption responds to income over and above the amount warranted by the PIH, i.e., the excess sensitivity of consumption to current income shocks. Some authors rationalize 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.[8] 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 behavior of consumers with short horizons, limited resources, or hyperbolic discount factors, giving an upper bound for the sensitivity of consumption to income shocks. [9]

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

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[7] Simulation results produced by Carroll (2001) 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.

[8] 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.

[9] An upper bound for the excess sensitivity parameter is ψ=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 behavior of households that do not use savings to buffer income shocks but spend all they receive.

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 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).[10] 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 labor market at age a_0 we can rewrite equation (3) as:

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

where $c_{i,a_0,t-a+a_0}$ reflects initial differences in preferences and endowments of individuals that belong to cohort *b*. Note that for notational 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:

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

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

[10] 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.

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 equations (2) and (4):

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

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

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.

In equation (5) the change in income inequality from one year to the next for a given cohort is due to the arrival of permanent and transitory shocks. In the absence of transitory shocks, income inequality unambiguously rises due to 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 behavior for such changes. In the PIH (ϕ=1, ψ=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 permanent and transitory shocks. Models with partial insurance ($0<\phi<1$, ψ=0) also predict a fanning out of cohort inequality, albeit at a slower pace than in the PIH. 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, i.e., the difference-in-difference of equations (5) and (6):

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

As we shall see, in the data we observe periods in which income and consumption inequality exhibit different trends, and equation (7) can be used to understand the forces behind this divergence. The third column of Table 2 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.[11] Thus, in models in which households smooth transitory income shocks, one needs an increase in income instability, i.e. $\Delta \operatorname{var}(e_{i,a,t}) > 0$, 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 paper we are particularly interested in estimating the impact of EMU on the excess sensitivity and insurance coefficients. We therefore estimate the above equation distinguishing between pre- and post-1999 observations:

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

$$
\begin{aligned}
&\Delta \operatorname{var}(\ln y_{i,a,t}) - \Delta \operatorname{var}(\ln c_{i,a,t}) = \\
&(1-\phi_N^2) \times (1-EMU) \times \operatorname{var}(u_{i,a,t}) + (1-\phi_E^2) \times EMU \times \operatorname{var}(u_{i,a,t}) + \\
&(1-\psi_N^2) \times (1-EMU) \times \operatorname{var}(e_{i,a,t}) + (1-\psi_E^2) \times EMU \times \operatorname{var}(e_{i,a,t}) - \operatorname{var}(e_{i,a,t-1})
\end{aligned}
\tag{8}
$$

Our test for the effect of financial integration is the joint test of the null hypothesis of no EMU effect, $(\phi_N = \phi_E)$ and $(\psi_N = \psi_E)$, against the hypothesis that both parameters have decreased after the introduction of the euro, either because they are more able to insure permanent shocks (as captured by a decrease in the insurance parameter ϕ), or because they are more able to smooth transitory shocks (as captured by a decrease in the excess sensitivity parameter ψ). As explained above, 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 ψ, in some specifications we impose the restriction $(\phi_N = \phi_E)$ and test only that the excess sensitivity coefficient has not changed over time $(\psi_N = \psi_E)$.

We run the regression based on equation (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 equation (8). Panel data on income allow us to identify the variances of income shocks, that is, the right-hand-side variables of equation (8). Omitting for simplicity the contribution of the observable characteristics X, we identify the cohort variances of income shocks non-parametrically using:

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

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

Note that identification of (9) requires four years of data on each household, while identification of (10) requires three years of data.

Therefore, 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 United States, where the CEX provides repeated cross-sectional data on consumption and the PSID provides panel data on income; or the United Kingdom, where repeated-cross section data on consumption data are available from the FES and panel data on income from the BHPS.

Finally, since the SHIW is conducted every two years (with a three-year gap in 1995-98) 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 equation (8).

5. Income and consumption inequality

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. For this purpose we use the Survey of Household Income and Wealth (SHIW), a representative sample of the Italian resident population conducted by the Bank of Italy.[12] 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

[12] The survey is available on line to all external users at www.bancaditalia.it. Questionnaire and documentation is available in English.

disposable income, distinguishing between after-tax earnings, transfers and income from capital.[13]

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 percent of the sample was re-interviewed in 1989, 27 percent in 1991, and about 45 percent after 1993.[14] Response rates in the panel section of the SHIW are generally above 70 percent, in line with other microeconomic data sets.[15] 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 ten (households interviewed each time from 1987 to 2006.

To minimize 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 unrecorded change in household head or measurement error. After these exclusions, the sample has about 50,000 consumption and income observations.

[13] Sampling is in two stages, first municipalities and then households. Municipalities are divided into 51 strata defined by 17 regions and 3 classes of population size (more than 40,000, 20,000 to 40,000, less than 20,000). Households are randomly selected from registry office records.
[14] In the panel component, the sampling procedure is also determined in two stages: (i) selection of municipalities (among those sampled in the previous survey); (ii) 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 4 times from 2000 to 2006) and a new component every survey (for instance, households re-interviewed only in 2006).
[15] For instance, the net response rate in the United States Consumer Expenditure Survey is slightly above 80 percent for the Interview and Diary samples.

5.1. The variance of income and consumption

Consumption is the sum of all expenditure categories except durables.[16] In our basic definition we also exclude rents and imputed rents, but in robustness checks we experiment with a broader definition of consumption and one that is adjusted for equivalence scales. Income is defined as the sum of labor 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 the permanent income hypothesis.[17]

Figure 3 reports the variance of log consumption and log disposable income from 1980 to 2006.[18] All statistics are computed using sample weights. Jappelli and Pistaferri (2008) 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 percent 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 percent between 1980 and 2006, after a 70 percent peak in 1998). The most plausible explanation for the increase in income inequality over the nineties points to extensive labor market reforms, raising labor market instability. Indeed, during the decade, fixed term contracts were deregulated, widening their use, temporary work agencies permitted, and restrictions concerning fixed term contracts for unskilled workers lifted. As a consequence, the overall index of Employment Protection Legislation (EPL) constructed by the OECD declined from 3.6 in the late 1980s, to 2.7 to in the late 1990s and 1.9 in 2003. According to this explanation, the increase in inequality is primarily of transitory nature (raising income instability).

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[16] Blanchard and Simon (2001) suggest that improvements in financial markets are associated with more non-durable consumption smoothing, but given the improved ability to borrow and lend may also "lead to a stronger stock-flow adjustment for purchases of durables, and thus potentially to more volatility of durable purchases" (p. 159). Thus, excluding durables from the definition of consumption is particularly important in our context.

[17] Adding back asset income or asset income net of imputed rents does not change the main results of the paper.

[18] 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.

Figure 3 also shows that inequality is lower for consumption than for income. Jappelli and Pistaferri (2008) find that the level of consumption inequality is higher for the definition of non-durable consumption net of housing rents. As for income, demographic variables absorb part of the income variability of income. Over time, consumption inequality increases in the last decade, but much less than income inequality: from 1980 to 2006 inequality increases by about 10 percent, after reaching a 20 percent peak in 2002).

Summing up, the increase in income inequality is not matched by a parallel increase in consumption inequality. Equation (9) 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 difference earnings process experienced by young and older 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.[19] Excluding individuals younger than 25 also 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

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

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 $\text{var}\ln(y_{i,a,t})$ and $\text{var}\ln(c_{i,a,t})$ of six cohorts born between 1936 and 1965 (notice that the cohort born in 1961-65 is observed only after 1985). The results confirm the three stylized facts emerging from the aggregate evidence in Figure 3: (1) cohort income inequality is substantially higher than consumption inequality; (2) for each cohort, there is a dramatic increase in income inequality in the early nineties (especially for the earlier cohorts), and a decline in the later part of the sample; (3) there is also an increase in consumption inequality for most cohorts, but the dynamics of consumption inequality is much smoother than that of income.

5.2. 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). This is the second important ingredient of our analysis. As explained in Section 3, we obtain non-parametric estimates of $\text{var}(e_{i,a,t})$ and $\text{var}(u_{i,a,t})$ using equations (9) and (10).

The two variances are displayed in Figure 5 for some of the cohorts of our constructed pseudo-panel. Over the sample period, we find 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 2 it is likely to derive from the labor market reforms and the associated greater labor flexibility. At the same time, there is no increase in the variance of permanent shocks (and even a decline, especially for the cohorts born in 1946-50 and 1951-55). The combined evidence suggests that the increase in income inequality that we observe in the early nineties is mainly attributable to an increase in the transitory component of inequality.

6. Empirical results

In this section we merge data on $\operatorname{var}\ln(y_{i,a,t})$, $\operatorname{var}\ln(c_{i,a,t})$ obtained from cross-sectional data with data on $\operatorname{var}(e_{i,a,t})$ and $\operatorname{var}(u_{i,a,t})$ obtained from the panel, and report estimates of the parameters of equation (8). In Table 2 we report structural estimates of the parameters of equation (8). Since each cohort is defined over one-year cells, the sample size includes 185 observations, corresponding to a maximum of 9 observations for each cohort; cells where the income shocks are computed on less than 5 observations. In the baseline specification we don't distinguish between pre- and post-EMU observations. The results reported in the first column show that the insurance parameter is ϕ=0.993 and the excess sensitivity parameter is ψ=0.322. Thus in our total sample estimate we find evidence that consumers don't smooth permanent shocks, and that consumption reacts also to transitory shocks, although much less then one-for-one.

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 percent of refunds within a quarter, and Parker that consumption reacts significantly to changes in tax rates. Jappelli and Pistaferri (2006), using data from the 1989-1995 SHIW, estimate the parameters that minimize 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 then predicted by the permanent income hypothesis. 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 interact the two variance shocks with the EMU dummy. The hypothesis that we test is that consumption has become less sensitive to income shocks after the introduction of the euro (a decrease in ϕ and ψ). The estimates indicate that both ϕ and ψ are indeed lower after the introduction of the euro. In particular, the insurance parameter falls from

1.087 to 0.727, and the excess sensitivity parameter from 0.310 to 0.285. The parameters are precisely estimated, and a formal statistical test of the joint hypothesis of no EMU effect rejects the null hypothesis that $(\phi_N = \phi_E)$ and $(\psi_N = \psi_E)$ at the 5 percent confidence level. In the third specification of Table 2 we constrain the insurance coefficient to be constant over time. The excess sensitivity coefficients are hardly affected, showing again a slight reduction after the introduction of the euro.

The last two columns repeat the estimation excluding cells where the income shocks variances are more reliable because are computed with at least 20 observations, reducing the sample size to 113 observations. The results only partly confirm the findings obtained for the total sample. The insurance coefficient falls from 0.967 to 0.839 after the introduction of the euro; the excess sensitivity coefficients are not statistically different from zero. In this case, 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. However, the hypothesis of no EMU effect is not rejected at conventional statistical levels. Restricting the sample to observations drawn from cells with at least 20 households does not change the picture.

A further experiment we perform is that we split the sample between households in which the head has completed high school and in which he of she has not, and then compute the variance of income shocks on the basis of cohorts defined over 5-years intervals.[20] This reduces the number of valid observations that we use to estimate the two parameters of equation (8), but allows estimation of different income processes for households with different levels of education.

Figure 6 reports income and consumption inequality from 1980 to 2006 for the two groups of households and three selected cohorts, and shows that that there is a much stronger increase in inequality among households with relatively low education. The estimates of separate income

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

processes for the two groups shows 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 high school. This is further evidence that the increase in income inequality in the last decade is due to the liberalization of the labor market, and the associated increase of 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 also regarding the structural parameters. In the basic specification the insurance parameter is close to one for both groups. The excess sensitivity parameter is higher among households with lower education (0.315) than among those who completed high school (0.121), suggesting that people with higher education have easier access to credit markets to smooth income fluctuations. When we distinguish between pre and post-EMU samples, we find that in both groups there is a considerable drop in the insurance parameter after the introduction of the euro, while the excess sensitivity parameter is imprecisely estimated in the high education group, both before and after the EMU policy shift. When we exclude cells with less then 20 households, the statistical test never supports the hypothesis that the EMU has increased consumption smoothing.

To check the robustness of the results, we perform several sensitivity checks: (1) 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; (2) on the other hand, since financial integration might impact consumption with a lag, we test the stability of the results we restrict the EMU sample to the years 2002-2006; (3) 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; (4) we subtract from consumption imputed rents, that are likely to be subject to substantial measurement error; (5) we deflate consumption and income by the OECD equivalence scale.[21] These experiments confirm the patterns found in Tables 2, 3 and 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.

[21] 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.

7. Conclusions

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 liberalization brought about by the introduction of the euro has made consumption less sensitive to income shocks in Italy. The paper 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 permanent income hypothesis (an insurance parameter close to one), although in some estimates 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 do decline after the introduction of the euro, but statistically the null hypothesis of no EMU effect is not rejected at standard confidence levels. 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 due to 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 on consumption smoothing opportunities. As highlighted by the European Commission (2008), financial integration remains 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 labor 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:

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

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

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

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

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

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

$$\begin{aligned}
\Delta^2 \operatorname{var}(\ln y_{i,a,t}) &= \operatorname{var}(u_{i,a,t}) + \operatorname{var}(u_{i,a-1,t-1}) + \Delta^2 \operatorname{var}(e_{i,a,t}) \\
\Delta^2 \operatorname{var}(\ln c_{i,a,t}) &= \phi^2 \left(\operatorname{var}(u_{i,a,t}) + \operatorname{var}(u_{i,a-1,t-1})\right) + \psi^2 \left(\operatorname{var}(e_{i,a,t}) + \operatorname{var}(e_{i,a-1,t-1})\right)
\end{aligned}$$

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

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

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

$$\begin{aligned}
E\left[(\ln y_{i,a,t} - \ln y_{i,a-2,t-2})(\ln y_{i,a+2,t+2} - \ln y_{i,a-4,t-4})\right] &= \operatorname{var}(u_{i,a,t}) + \operatorname{var}(u_{i,a-1,t-1}) \\
-E\left[(\ln y_{i,a,t} - \ln y_{i,a-2,t-2})(\ln y_{i,a+2,t+2} - \ln y_{i,a,t})\right] &= \operatorname{var}(e_{i,a,t}) \\
-E\left[(\ln y_{i,a-2,t-2} - \ln y_{i,a-4,t-4})(\ln y_{i,a,t} - \ln y_{i,a-2,t-2})\right] &= \operatorname{var}(e_{i,a-2,t-2})
\end{aligned}$$

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

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

and hence rewrite (A.5) as:

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

This is the regression we run and whose results are reported in Tables 2-4. We use a similar strategy to deal with the 3-year gap between the 1995 and 1998 surveys.

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Table 1
Implications of various models for the change in consumption inequality

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

Table 2
Difference-in-difference of var(y) and var(c), one-year cohorts

	Total sample	Total sample	Total sample	Excluding n<20	Excluding n<20	Excluding n<20
ϕ	0.993		0.989	0.913		0.912
	(0.052)**		(0.052)**	(0.091)**		(0.093)**
ψ	0.322			0.079		
	(0.065)**			(0.346)		
ϕ pre-EMU		1.087			0.967	
		(0.056)**			(0.119)**	
ϕ post-EMU		0.727			0.839	
		(0.125)**			(0.151)**	
ψ pre-EMU		0.310	0.349		-0.054	0.064
		(0.084)**	(0.075)**		(0.603)	(0.504)
ψ post-EMU		0.285	0.270		0.192	0.107
		(0.119)*	(0.128)*		(0.270)	(0.450)
Observations	185	185	185	113	113	113
P-value		0.031	0.594		0.787	0.945

Note. The table reports structural estimates of various versions of equation (8) in the text. The p-value is the probability value of the test that ϕ and ψ have not changed after the introduction of the euro. Robust standard errors are reported in parenthesis.

Table 3
Difference-in-difference of var(y) and var(c), three-years cohorts

	Total sample			Excluding n<20		
ϕ	1.106		1.104	0.794		0.794
	(0.035)**		(0.035)**	(0.110)**		(0.111)**
ψ	0.240			0.306		
	(0.113)*			(0.080)**		
ϕ pre-EMU		1.132			0.885	
		(0.034)**			(0.131)**	
ϕ post-EMU		0.665			0.649	
		(0.207)**			(0.209)**	
ψ pre-EMU		0.244	0.270		0.278	0.307
		(0.126)	(0.119)*		(0.111)*	(0.097)**
ψ post-EMU		0.290	0.142		0.326	0.302
		(0.167)	(0.347)		(0.127)*	(0.135)*
Observations	87	87	87	56	56	56
		0.087	0.726		0.634	0.975

Note. The table reports structural estimates of various versions of equation (8) in the text. The p-value is the probability value of the test that ϕ and ψ have not changed after the introduction of the euro. Robust standard errors are reported in parenthesis.

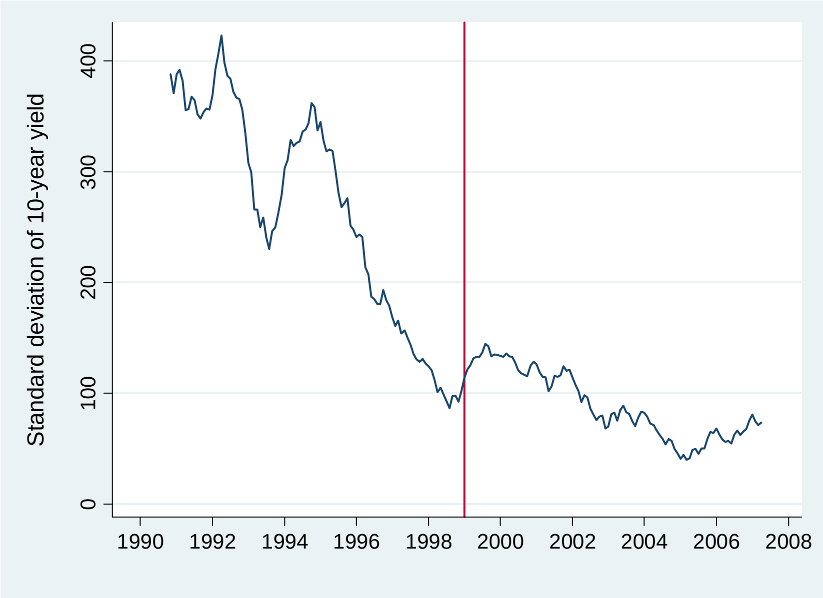
Table 4
Difference-in-difference of var(y) and var(c), 5-year cohorts by education

Low education	Total sample			Excluding n<20		
ϕ	0.997		0.994	0.990		1.012
	(0.053)**		(0.053)**	(0.091)**		(0.090)**
ψ	0.315			0.293		
	(0.137)*			(0.112)*		
ϕ pre-EMU		1.022			1.048	
		(0.053)**			(0.110)**	
ϕ post-EMU		0.595			0.943	
		(0.308)			(0.165)**	
ψ pre-EMU		0.340	0.362		0.335	0.345
		(0.139)*	(0.133)**		(0.110)**	(0.103)**
ψ post-EMU		0.266	-0.172		-0.124	-0.225
		(0.362)	(0.550)		(0.604)	(0.298)
Observations	55	55	55	35	35	35
P-value		0.332	0.346		0.467	0.072

High education	Total sample			Excluding n<20		
ϕ	1.043		1.043	0.771		0.740
	(0.060)**		(0.060)**	(0.121)**		(0.132)**
ψ	0.121			0.350		
	(0.285)			(0.089)**		
ϕ pre-EMU		1.106			0.853	
		(0.057)**			(0.145)**	
ϕ post-EMU		0.517			0.511	
		(0.302)			(0.305)	
ψ pre-EMU		-0.255	-0.254		0.410	0.448
		(0.192)	(0.206)		(0.138)**	(0.123)**
ψ post-EMU		0.265	0.273		0.284	0.289
		(0.161)	(0.167)		(0.135)*	(0.133)*
Observations	53	53	53	32	32	32
P-value		0.024	0.053		0.460	0.389

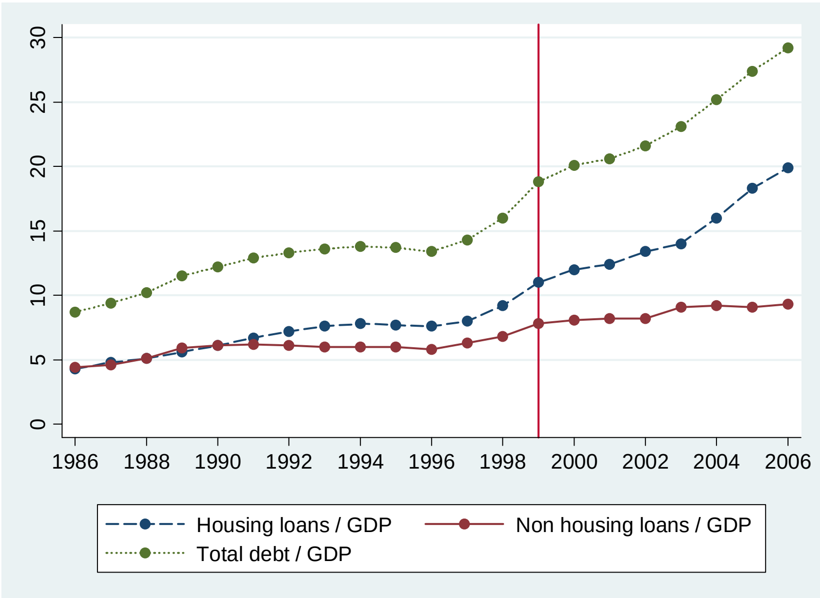
Note. The table reports structural estimates of various versions of equation (8) in the text. The p-value is the probability value of the test that ϕ and ψ have not changed after the introduction of the euro. Robust standard errors are reported in parenthesis.

Figure 1
Standard deviation of the 10-year benchmark bond yield spreads in the EMU countries



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) in the EMU countries (Austria, Belgium, Finland, France, Greece, Ireland, Italy, Netherlands, Portugal, Spain). Source: Datastream.

Figure 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.

Figure 3
Income and consumption inequality, 1980-2006

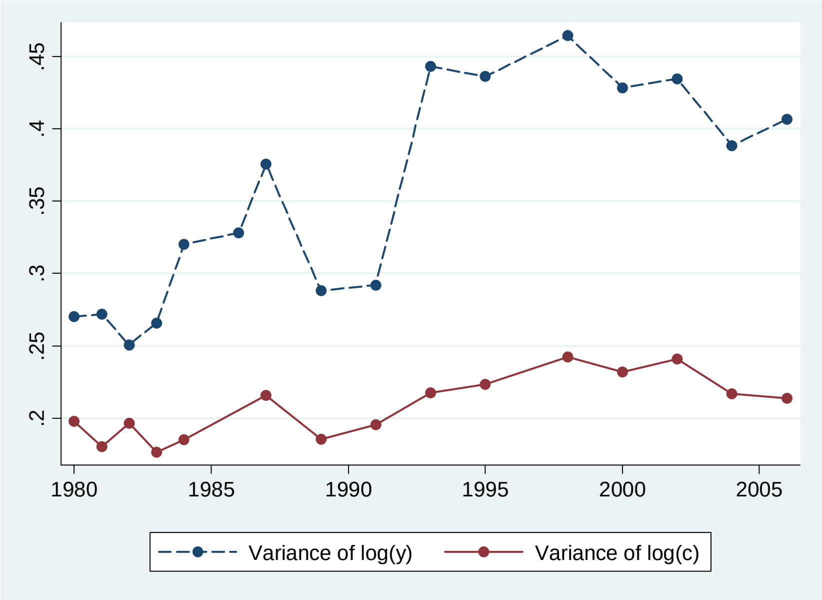


Figure 4
Consumption and income inequality by selected cohorts

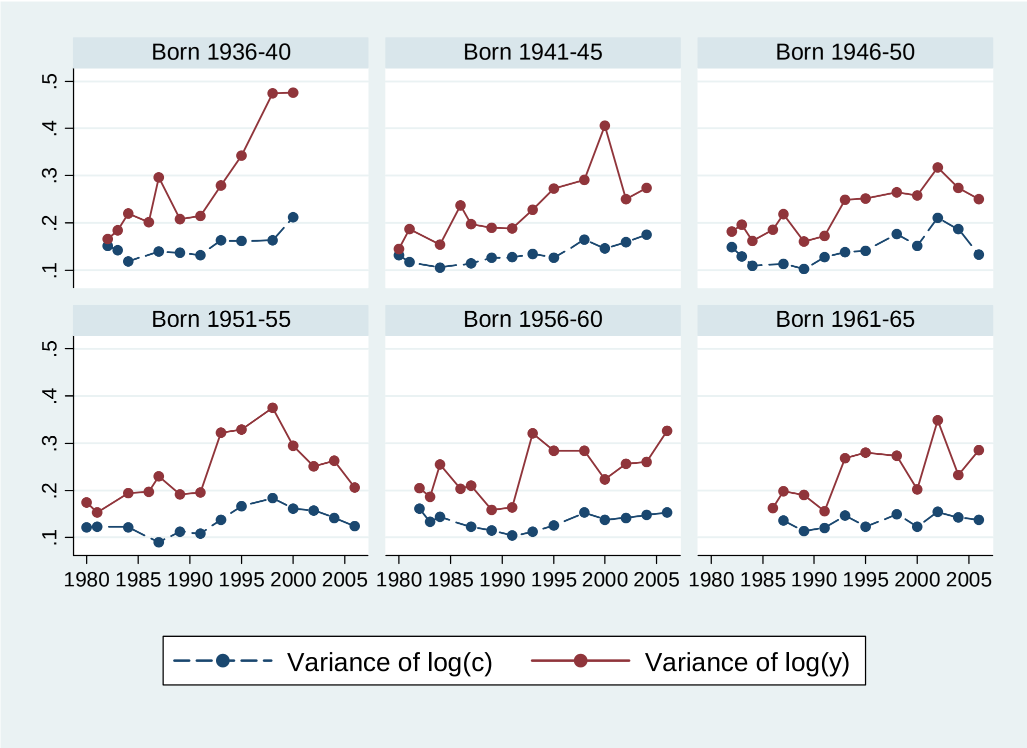


Figure 5
The variance of transitory and permanent income shocks by selected cohorts

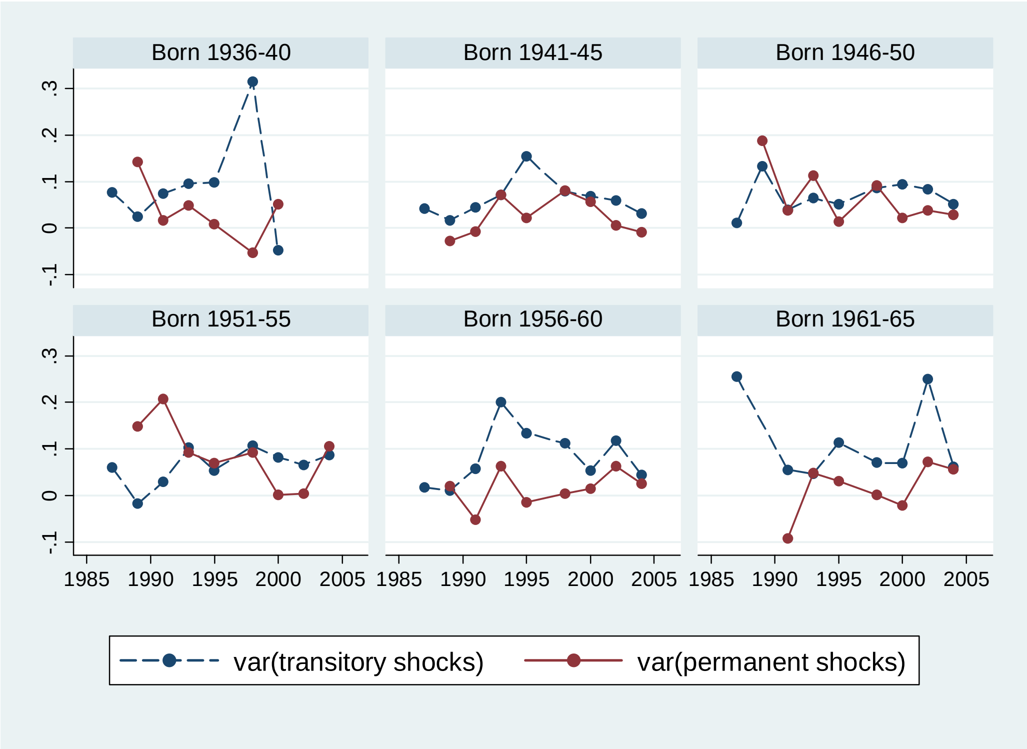


Figure 6
Consumption and income inequality by selected cohorts and education groups