Financial Integration and Consumption Smoothing

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Abstract

We present a new empirical strategy for testing if financial integration improves household consumption smoothing. Our test uses microeconomic data and 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 find that the process of financial market integration and liberalization brought about by the introduction of the euro has not affected the sensitivity of consumption with respect to income shocks in Italy. The paper makes a significant contribution also 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.

Keywords: Risk Sharing, Consumption Smoothing, Financial Market Integration.

JEL codes: D91

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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 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, 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.
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. 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. 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 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 behavior of consumption and income. Finally, to distinguish more sharply the EMU effect from potential confounds, 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 U.K. (a country that didn’t join). For this robustness check, we draw consumption and income data from the U.K. 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 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 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 6 presents the baseline results, and Section 7 the results of the robustness analysis. Section 8 concludes.
2. 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 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

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1 See Jappelli and Pagano (2008) for a survey of the real effects of financial market integration in the context of the EMU.
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 (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 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 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, disability, etc.). In this paper 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.

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.
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 (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.\(^4\)

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 (2010) 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.

\(^3\) 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).

\(^4\) Recent microeconomic evidence also throws light on the role that international financial integration can play in improving the allocation of capital across firms. Galindo, 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.
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. In Italy the convergence toward zero of the spread over the German yield is dramatic, as shown in Figure 1. The figure also plots the spread for the U.K., 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 U.K., which was pursuing an independent monetary policy. The U.K. therefore represents an interesting country to compare with Italy, an issue that we will take up econometrically in Section 7.

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 inefficiency in Italy are higher than in countries at a similar level of financial development. Casolaro, Gambacorta and Guiso (2006) 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

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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 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.

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.
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. 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.

4. The empirical strategy

We rely on the covariance restriction implied by the permanent income hypothesis 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 (see Deaton, 1991; Carroll, 1997; and 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):

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

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. We comment on the plausibility of this income process in Section 5.2.

In this paper 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. Recent papers have explored the implications of endogenizing income for consumption allocations through human capital accumulation, job search, labor supply and cross-firm mobility. Accounting for endogenous income, however, is beyond the purpose of the present study.9

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 \(\pi\) 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:

\[
\text{var}_b \left( \ln y_{i,a,t} \right) = \text{var}_b \left( \pi_{i,a_0,t-a+a_0} \right) + \sum_{j=a_0+1}^{a} \text{var}_b \left( u_{i,j,t-a+j} \right) + \text{var}_b \left( e_{i,a,t} \right)
\]

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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, Meghir and Pistaferri (2010) and Heathcote, Storesletten and Violante (2009) and Postel-Vinay and Thuron (2009) for recent applications.
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 $\pi$, $u$, 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:

$$
\ln c_{i,a,d,t} \approx \ln c_{i,a-1,t-1} + z_{i,a,d} \times \gamma + \phi u_{i,a,d} + \psi e_{i,a,d}
$$

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 ($\phi=1$), while transitory shocks have negligible effects ($\psi=0$) because consumers use assets to smooth temporary income fluctuations. The buffer stock model delivers similar implications. 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 ($\psi=0$ and $0<\phi<1$).

The parameter $\psi$ 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 transitory income shocks. Some authors rationalize excess sensitivity by

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10 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. 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.
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. 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.

Consistent with the models’ predictions, we denote $\psi$ 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 $\phi$ 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 a reduction over time in the sensitivity of consumption to transitory shocks ($\psi$). The effect of financial market integration on the sensitivity to permanent shocks ($\phi$) 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). For

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11 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 $i-1$ chooses assets $a_{i-1}$ to constrain the consumption of self $i$. 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.

12 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 ($\ln c_{a,i} = \ln y_{a,i}$), 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.

13 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.
these reasons, one should expect that financial integration might impact consumption primarily through a change in $\psi$ rather than in $\phi$.

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:

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

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 equations (2) and (4):

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

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

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 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 \((\phi=1, \psi=0)\), 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. 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):

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

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 than 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.\textsuperscript{14} Thus, in models in which households smooth transitory income shocks, one needs an increase in income instability, i.e. $\Delta \text{var}(e_{i,a,t}) > 0$, 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 paper 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 equation (7) allowing the sensitivity of consumption to income shocks to be affected linearly by the introduction of the euro:

$$
\Delta \text{var}(\ln y_{i,a,t}) - \Delta \text{var}(\ln c_{i,a,t}) = \left(1 - (\phi + \phi_E E)^2\right) \text{var}(u_{i,a,t}) \\
+ \left(1 - (\psi + \psi_E E)^2\right) \text{var}(e_{i,a,t}) - \text{var}(e_{i,a,t-1})
$$

We test for the effect of financial integration by looking at the p-value of the joint null hypothesis of no EMU effect, $(\phi_E = \psi_E = 0)$, 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. In the robustness analysis we study if alternative measures of financial integration, such as the interest rate spread, affect the estimates of the coefficients of interest. 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 the sensitivity of consumption to transitory shocks, in some specifications we impose the restriction $\phi_E = 0$ and test only that sensitivity of consumption to transitory income shocks has not changed over time.

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

\textsuperscript{14} 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 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 $X$, as in Meghir and Pistaferri (2004), we identify the cohort-specific variances of income shocks non-parametrically using:

$$E[(\ln y_{i,t} - \ln y_{i,t-1})(\ln y_{i,t+1} - \ln y_{i,t-2})] = \text{var}(u_{i,t})$$  \hspace{1cm} (9)

$$-E[(\ln y_{i,t} - \ln y_{i,t-1})(\ln y_{i,t+1} - \ln y_{i,t})] = \text{var}(e_{i,t})$$  \hspace{1cm} (10)

Note that identification of (9) requires four years of data on each household, while identification of (10) requires three years of data. Furthermore, if income is measured with classical i.i.d. error, one can prove that equation (9) still identifies the variance of permanent shocks, while equation (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 $\phi$ is unbiased, while the estimate of $\psi$ is upward biased.\textsuperscript{15} Clearly, unless measurement error changes systematically over time, there is no reason to believe that our test of financial integration is affected by 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 (equation 6).

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. As we shall see in Section 7, another application is the United Kingdom, where repeated-cross section data on consumption and income data are available from the FES and panel data on income from the BHPS.

\textsuperscript{15} In particular, $\rho \lim \bar{\psi} = \sqrt{1 - \frac{\lambda}{1 - \psi^2}}$ where $\lambda$ is the signal-to-noise ratio, i.e., the ratio of the variance of transitory shocks to the sum of the variance of transitory shocks and the variance of the measurement error.
5. 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 Survey of Household Income and Wealth (SHIW), a representative sample of the Italian resident population conducted by the Bank of Italy.\(^{16}\) 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.\(^{17}\)

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.\(^{18}\)

Response rates in the panel section of the SHIW are generally above 70 percent, in line with

\(^{16}\) The survey is available on line to all external users at [www.bancaditalia.it](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.

\(^{17}\) 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.

\(^{18}\) 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).
other microeconomic data sets. 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.

5.1. The variance of income and consumption

Consumption is the sum of all expenditure categories except durables. 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 intertemporal consumption decisions.

Figure 3 reports the variance of log consumption and log disposable income from 1980 to 2006. 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

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19 For instance, the net response rate in the United States Consumer Expenditure Survey is slightly above 80 percent for the Interview and Diary samples.
20 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.
21 In Section 7 we check the robustness of the results using alternative definitions of income.
22 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.
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).

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. 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 (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 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. 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 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 $\ln(\text{var}_{i,a,t})$ and $\ln(\text{var}_{i,a,t})$ 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. 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), which assume that income is the sum of a random walk permanent component and a serially uncorrelated transitory component.

\footnote{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.}
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):

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

Pooling data for all years, we find that the estimated autocovariance at the first order \((j=1)\) is \(-0.0798\) (s.e. \(0.0050\)), but autocovariances after the first order drop abruptly to zero and are statistically insignificant.\(^{24}\) 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 1995-98) 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 \(\text{var}(e_{i,a,t})\) can be approximated by a smooth function of adjacent variances, and reformulate accordingly the estimated equations (7) and (8).\(^{25}\)

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 5.1 it is likely to derive from the labor market reforms and the associated greater labor 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

\(^{24}\)For example, the 2\(^{nd}\) order autocovariance is \(-0.0044\) (s.e. \(0.0040\)) and the 3\(^{rd}\) order autocovariance is \(-0.0016\) (0.0045).

\(^{25}\)We take into account the two-year gap issue when estimating the income autocovariances discussed above.
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 \(\text{var}(\ln(y_{i,a,t}))\), \(\text{var}(\ln(c_{i,a,t}))\) obtained from cross-sectional data with data on \(\text{var}(e_{i,a,t})\) and \(\text{var}(u_{i,a,t})\) 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). These are obtained by non-linear least squares correcting the standard errors for heteroskedasticity of unknown form. 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 are dropped. 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 \(\phi=0.989\) and the excess sensitivity parameter is \(\psi=0.282\). 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 than one-for-one.\(^{26}\)

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

\(^{26}\) Recall from footnote 15 that in the presence of measurement error in income parameter \(\psi\) is upward biased. The estimate in column 1 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.
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 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 $\phi_E$ and $\psi_E$). The estimates indicate
that $\phi_E$ is negative (-0.378) and statistically different from zero at the 5 percent level after the
introduction of the euro, while $\psi_E$ 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 percent
confidence level. In the third specification of Table 2 we constrain the insurance coefficient to be
constant over time ($\phi_E=0$). 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 excluding cells where the income shocks
variances are more reliable because 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 statically different from zero at the one percent
level ($\phi=0.947$), while 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 $\psi_E$ 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 that we split the sample between households in which the head has completed college and in which he or she has not, and then compute the variance of income shocks on the basis of cohorts defined over 5-years intervals.27 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 5 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 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 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. 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.

### 7. 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 behavior of consumption and income. We also compare statistically the dynamics of

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27 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.
consumption and income inequality in Italy with the dynamics of the same variables in the U.K. All the regressions in this Section use cells defined on the basis of one-year cohorts.

7.1. Alternative measure of financial integration

Recall that in equation (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 an alternative measure of financial integration:

$$
\Delta \text{var}(\ln y_{t,i,a,d}) - \Delta \text{var}(\ln c_{t,i,a,d}) = \left(1 - (\phi + \phi_E S)^2\right) \text{var}(u_{t,i,a,d}) 
+ \left(1 - (\psi + \psi_E S)^2\right) \text{var}(e_{t,i,a,d}) - \text{var}(e_{t,i,a,d-1})
$$

(11)

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 $\psi_E > 0$. 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 $\phi_E$ and $\psi_E$ are positive, but not statistically different from zero, confirming the results in Section 6.

In a related test, we split the sample between regions that are more and less financially backward according to the indicator developed in Guiso, Sapienza and Zingales (2004). 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.

7.2. 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 private transfers. To check whether our results are sensitive to this criticism, in the third and fourth column of Table 4 we estimate the income process and compute the income shock variances 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 than decreased) after the introduction of the euro.

7.3. 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 (equation 1) and log consumption (equation 3) is:

\[
\text{cov}(\ln c_{i,a,t}, \ln y_{i,a,t}) = \text{var}(\ln c_{i,a_0,t-a+a_0}, \ln y_{i,a_0,t-a+a_0}) + \phi \sum_{j=a_0}^{a} \text{var}(u_{i,j,t-a+j}) + \psi \text{var}(e_{i,a,t})
\]

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

\[
\Delta \text{cov}(\ln c_{i,a,t}, \ln y_{i,a,t}) = \phi \text{var}(u_{i,a,t}) + \psi \Delta \text{var}(e_{i,a,t})
\]

We therefore estimate jointly equations (8) and (12) imposing cross-equations 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 \(\phi\) and \(\psi\) have statistically significantly decreased after the introduction of the Euro.
7.4. Using U.K. cohorts as a control group

Our identification strategy relies on the assumption that the only reason why $\phi$ and $\psi$ 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 $\phi$ and $\psi$ over time might result from offsetting effects). To account for this possibility we expand the analysis considering the U.K. 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 U.K. 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 U.K. 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 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.28 Figure 6 is the equivalent of Figure 4 for the U.K. 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 equation (8) which allows for post-1999 effect common to both countries ($\phi_E$ and $\psi_E$), and a post-1999 effect specific only to Italy ($\phi_{ITALY}$ and $\psi_{ITALY}$):29

$$\Delta \text{var}(\ln y_{i,a,t}) - \Delta \text{var}(\ln c_{i,a,t}) = \left(1 - \left(\phi + \phi_E E + \phi_{ITALY} ITALY \times E \right)^2\right) \text{var}(u_{i,a,t}) $$

$$+ \left(1 - \left(\psi + \psi_E E + \psi_{ITALY} ITALY \times E \right)^2\right) \text{var}(e_{i,a,t}) - \text{var}(e_{i,a,t-1})$$

(13)

28 A full description of the two U.K. datasets and sample selection is contained in Blundell and Etheridge (2010).
29 We experiment also with a less parsimonious specification including the terms $\phi_{ITALY}$ and $\psi_{ITALY}$, 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.
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 U.K. after 1999 (if anything, the sensitivity with respect to permanent shocks increases in the later part of the sample in both countries).

7.5. Additional tests

To further check the robustness of the results, we perform a number of additional 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 2002-06; (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. 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.

8. Conclusions

In this paper 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.

30 The OECD equivalence scale is defined as \( E = 1 + 0.5 \times (\text{number of children}) + 0.7 \times (\text{number of adult members} - 1) \). A child is any household member aged 16 or less.
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 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 do 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 U.K., and exploit the additional restrictions imposed by the theory on the covariance between income and consumption. All 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 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:

\[ \text{var}(\ln y_{i,a,t}) = \text{var}(\pi_{i,a_0,t-a+\alpha_0}) + \sum_{j=a_0+1}^a \text{var}(u_{i,j,t-a+j}) + \text{var}(e_{i,a,t}) \quad \text{(A.1)} \]

\[ \text{var}(\ln c_{i,a,t}) = \text{var}(\ln c_{i,a_0,t-a+\alpha_0}) + \phi^2 \sum_{j=a_0+1}^a \text{var}(u_{i,j,t-a+j}) + \psi^2 \sum_{j=a_0+1}^a \text{var}(e_{i,j,t-a+j}) \quad \text{(A.2)} \]

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

\[ \text{var}(\ln y_{i,a-2,t-2}) = \text{var}(\pi_{i,a_0,t-a+\alpha_0}) + \sum_{j=a_0+1}^{a-2} \text{var}(u_{i,j,t-a+j}) + \text{var}(e_{i,a-2,t-2}) \quad \text{(A.3)} \]

\[ \text{var}(\ln c_{i,a-2,t-2}) = \text{var}(\ln c_{i,a_0,t-a+\alpha_0}) + \phi^2 \sum_{j=a_0+1}^{a-2} \text{var}(u_{i,j,t-a+j}) + \psi^2 \sum_{j=a_0+1}^{a-2} \text{var}(e_{i,j,t-a+j}) \quad \text{(A.4)} \]

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

\[ \Delta^2 \text{var}(\ln y_{i,a,t}) = \text{var}(u_{i,a,t}) + \Delta^2 \text{var}(e_{i,a,t}) \]

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

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

\[ \Delta^2 \text{var}(\ln y_{i,a,t}) - \Delta^2 \text{var}(\ln c_{i,a,t}) = (1 - \phi^2) (\text{var}(u_{i,a,t}) + \text{var}(u_{i,a-1,t-1})) + (1 - \psi^2) \text{var}(e_{i,a,t}) - \psi^2 \text{var}(e_{i,a-1,t-1}) - \text{var}(e_{i,a-2,t-2}) \quad \text{(A.5)} \]

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

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

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

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

However, \( \text{var}(e_{i,a-1,t-1}) \) remains not identified. We assume that it can be approximated by a smooth function of adjacent variances. In this specific case, we assume:
\[ \text{var}\left(e_{i,a-1,j-1}\right) = \frac{\text{var}\left(e_{i,a,j}\right) + \text{var}\left(e_{i,a-2,j-2}\right)}{2} \]

and hence rewrite (A.5) as:

\[
\Delta^2 \text{var}\left(\ln y_{i,a,t}\right) - \Delta^2 \text{var}\left(\ln c_{i,a,t}\right) = \left(1 - \phi^2\right)\left(\text{var}\left(u_{i,a,t}\right) + \text{var}\left(u_{i,a-1,j-1}\right)\right) \\
+ \left(1 - \psi^2 - \frac{\psi^2}{2}\right)\text{var}\left(e_{i,a,d}\right) - \left(1 + \frac{\psi^2}{2}\right)\text{var}\left(e_{i,a-2,j-2}\right)
\]

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.
References


European Commission (2008), “EMU@10: Successes and Challenges after 10 years of Economic and Monetary Union,” European Economy 2.


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

<table>
<thead>
<tr>
<th>Model</th>
<th>Change in the variance of consumption (\Delta \text{var}\ln(c_{i,a,t}))</th>
<th>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}))</th>
<th>Restrictions</th>
</tr>
</thead>
<tbody>
<tr>
<td>PIH</td>
<td>(\text{var}(u_{i,a,t}))</td>
<td>(\Delta \text{var}(e_{i,a,t}))</td>
<td>(\phi = 1, \psi = 0)</td>
</tr>
<tr>
<td>Partial Insurance</td>
<td>(\phi^2 \text{var}(u_{i,a,t}))</td>
<td>((1 - \phi^2) \text{var}(u_{i,a,t}) + \Delta \text{var}(e_{i,a,t}))</td>
<td>(0 &lt; \phi &lt; 1, \psi = 0)</td>
</tr>
<tr>
<td>Complete Markets</td>
<td>0</td>
<td>(\text{var}(u_{i,a,t}) + \Delta \text{var}(e_{i,a,t}))</td>
<td>(\phi = \psi = 0)</td>
</tr>
<tr>
<td>Excess sensitivity</td>
<td>(\text{var}(u_{i,a,t}) + \psi^2 \text{var}(e_{i,a,t}))</td>
<td>(((1 - \psi^2) \text{var}(e_{i,a,t}) - \text{var}(e_{i,a,t-1}))</td>
<td>(\phi = 1, 0 &lt; \psi \leq 1)</td>
</tr>
</tbody>
</table>
### Table 2
**Difference-in-difference of var(y) and var(c), one-year cohorts**

<table>
<thead>
<tr>
<th></th>
<th>Total sample</th>
<th>Excluding n&lt;30</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>φ</strong></td>
<td>0.989</td>
<td>0.861</td>
</tr>
<tr>
<td></td>
<td>(0.056)**</td>
<td>(0.114)**</td>
</tr>
<tr>
<td><strong>ψ</strong></td>
<td>0.282</td>
<td>0.224</td>
</tr>
<tr>
<td></td>
<td>(0.081)**</td>
<td>(0.139)</td>
</tr>
<tr>
<td><strong>φ_E</strong></td>
<td>-0.378</td>
<td>-0.271</td>
</tr>
<tr>
<td></td>
<td>(0.152)*</td>
<td>(0.266)</td>
</tr>
<tr>
<td><strong>ψ_E</strong></td>
<td>0.030</td>
<td>0.374</td>
</tr>
<tr>
<td></td>
<td>(0.178)</td>
<td>(0.705)</td>
</tr>
<tr>
<td>Observations</td>
<td>185</td>
<td>82</td>
</tr>
</tbody>
</table>

Note. 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. One asterisk denotes significance at 1% level; two asterisks significance at the 5% level.

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

<table>
<thead>
<tr>
<th></th>
<th>Total sample</th>
<th>Excluding n&lt;30</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>φ</strong></td>
<td>1.103</td>
<td>0.801</td>
</tr>
<tr>
<td></td>
<td>(0.035)**</td>
<td>(0.118)**</td>
</tr>
<tr>
<td><strong>ψ</strong></td>
<td>0.205</td>
<td>0.273</td>
</tr>
<tr>
<td></td>
<td>(0.133)</td>
<td>(0.095)**</td>
</tr>
<tr>
<td><strong>φ_E</strong></td>
<td>-0.481</td>
<td>-0.199</td>
</tr>
<tr>
<td></td>
<td>(0.216)*</td>
<td>(0.273)</td>
</tr>
<tr>
<td><strong>ψ_E</strong></td>
<td>0.121</td>
<td>0.107</td>
</tr>
<tr>
<td></td>
<td>(0.232)</td>
<td>(0.196)</td>
</tr>
<tr>
<td>Observations</td>
<td>87</td>
<td>53</td>
</tr>
</tbody>
</table>

Note. 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. One asterisk denotes significance at 1% level; two asterisks significance at the 5% level.
Table 4  
Difference-in-difference of \( \text{var}(y) \) and \( \text{var}(c) \), 5-year cohorts by education

<table>
<thead>
<tr>
<th></th>
<th>No college education</th>
<th></th>
<th>College education</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>( \phi )</td>
<td>0.893 (0.125)**</td>
<td>0.900 (0.171)**</td>
<td>0.910 (0.127)**</td>
<td>0.720 (0.151)**</td>
</tr>
<tr>
<td>( \psi )</td>
<td>0.327 (0.109)**</td>
<td>0.356 (0.115)**</td>
<td>0.354 (0.110)**</td>
<td>0.337 (0.109)**</td>
</tr>
<tr>
<td>( \phi_E )</td>
<td>0.023 (0.261)</td>
<td></td>
<td>-0.221 (0.348)</td>
<td></td>
</tr>
<tr>
<td>( \psi_E )</td>
<td>-0.209 (0.568)</td>
<td>-0.187 (0.442)</td>
<td>-0.160 (0.198)</td>
<td>-0.186 (0.185)</td>
</tr>
<tr>
<td>Observations</td>
<td>32 32 32</td>
<td>31 31 31</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note. The table reports non-linear least squares estimates of various versions of equation (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. One asterisk denotes significance at 1% level; two asterisks significance at the 5% level. Robust standard errors are reported in parenthesis.
### Table 5
**Robustness analysis**

<table>
<thead>
<tr>
<th></th>
<th>Using the spread as a measure of integration</th>
<th>Using family earnings</th>
<th>Using covariance restrictions</th>
<th>Using UK as a comparison</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Total sample</td>
<td>Excluding n&lt;30</td>
<td>Total sample</td>
<td>Excluding n&lt;30</td>
</tr>
<tr>
<td>( \phi )</td>
<td>0.639</td>
<td>0.921</td>
<td>0.899</td>
<td>0.666</td>
</tr>
<tr>
<td></td>
<td>(0.176)**</td>
<td>(0.284)**</td>
<td>(0.075)**</td>
<td>(0.248)**</td>
</tr>
<tr>
<td>( \psi )</td>
<td>0.033</td>
<td>-0.009</td>
<td>-0.130</td>
<td>0.045</td>
</tr>
<tr>
<td></td>
<td>(0.016)*</td>
<td>(0.041)</td>
<td>(0.153)</td>
<td>(0.311)</td>
</tr>
<tr>
<td>( \phi_E )</td>
<td>0.028</td>
<td>0.006</td>
<td>0.263</td>
<td>0.339</td>
</tr>
<tr>
<td></td>
<td>(0.282)</td>
<td>(0.412)</td>
<td>(0.120)*</td>
<td>(0.150)*</td>
</tr>
<tr>
<td>( \psi_E )</td>
<td>0.027</td>
<td>0.029</td>
<td>0.250</td>
<td>0.165</td>
</tr>
<tr>
<td></td>
<td>(0.025)</td>
<td>(0.041)</td>
<td>(0.146)</td>
<td>(0.171)</td>
</tr>
<tr>
<td>( \phi\text{ITA,E} )</td>
<td>-0.164</td>
<td>-0.581</td>
<td>-0.224</td>
<td>-0.321</td>
</tr>
<tr>
<td>( \psi\text{ITA,E} )</td>
<td>-0.027</td>
<td>-0.070</td>
<td>(0.246)</td>
<td>(0.309)</td>
</tr>
<tr>
<td>Observations</td>
<td>185</td>
<td>82</td>
<td>180</td>
<td>61</td>
</tr>
</tbody>
</table>

Note. 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 equation (11) in the text. In columns (3) and (4) we use family earnings as income measure. In columns (5) and (6) we use equations (8) and (12) jointly to estimate the consumption parameters. In the last two columns we estimate equation (13) in the text, using observations from both Italy and the U.K. “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. One asterisk denotes significance at 1% level; two asterisks significance at the 5% level.
Figure 1
10-year benchmark bond yield spread in Italy and the U.K.

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.
Figure 2
Household debt-GDP in Italy

Figure 3
Income and consumption inequality, 1980-2006
Figure 4
Consumption and income inequality by selected cohorts
Figure 5
Consumption and income inequality by selected cohorts and education groups
Figure 6
Consumption and income inequality by selected cohorts in the U.K.