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Abstract

This paper examines whether or not consumption risk sharing occurs in a panel of industrialized countries. We theoretically derive the international consumption insurance proposition in a simple setup and show how it should be modified in more complicated models. We empirically analyze the implications of the proposition for pairs of countries over cycles of different length and find that (i) aggregate domestic consumption is completely insured against idiosyncratic real, demographic, fiscal and monetary shocks, but that it covaries with domestic variables over long or infinite cycles. Also, the cross equation restrictions imposed by the theory are, in general, rejected. The policy implications of the results are discussed.

JEL Classification No.: D81, E32, F21.

Key words: Consumption Insurance, International Financial Markets, Long Run Convergence.

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1 Introduction

The idea that agents attempt to insure their consumption streams against individual income or wealth fluctuations is a pervasive feature of many modern dynamic macro models. It arises in closed economy setups where there is heterogeneity in income or preferences across agents (see e.g. Mace (1991) or Marcet and Marimon (1992)), in open economy setups where countries with heterogeneous income streams borrow and lend internationally in order to bear only aggregate world–wide risk (see e.g. Backus, Kehoe and Kydland (1992)), and in models where consumption insurance may be a trigger (or a deterrent) for long term growth (see e.g. Devereux and Smith (1991) and Obstfeld (1992b)). As emphasized by Cochrane (1991), the basic idea of consumption insurance is in some respects the cross–sectional counterpart of the permanent income hypothesis. Complete insurance implies that the consumption of individual economic units (agents, families or countries) should not vary in response to idiosyncratic shocks while the permanent income hypothesis implies that the consumption of an individual economic unit should not vary over time in response to idiosyncratic transitory shocks.

Full consumption insurance obtains theoretically when financial markets are complete or when there is a set of institutions which implement Pareto optimal allocations. Some of these institutional arrangements do exist in the real world. At the individual agent level, unemployment or medical insurance schemes, welfare and social government programs or even intergenerational transfers may help in reducing the effect of individual specific shocks. At a country level, charities or disaster relief programs, automatic international lending agreements or direct foreign aid may help to insure national consumption from catastrophic, idiosyncratic income fluctuations. Completeness of financial markets or the presence of institutions implementing first best allocations are, however, only sufficient conditions for full consumption insurance to take place. For example, Duffie and Huang (1985) show that continuous trading of a few long lived securities may implement Pareto optimal allocations. Similarly, recent work by Baxter and Crucini (1992) and Marcet and Singleton (1992) show that close to full insurance obtains even in situations where financial markets are strongly incomplete and no mechanism for implementing first best solutions exists as long as agents have similar preferences (like more consumption to less and less variability to more) and differ only in their income streams.

Although the mechanics of consumption insurance are very simple theoretically, casual empiricism suggests that international risk sharing is hardly a feature of the real world. It is often claimed that aggregate consumption of individual countries *do* react to country specific shocks. In support of this idea it is typically reported that poor underdeveloped countries starve when droughts or civil wars cut their own food supplies. For these countries, insurance markets seem to be imperfect or at least at least partially inaccessible, no institutions implementing first best exist and borrowing opportunities are limited (see Atkeson (1991) or Atkeson and Lucas (1992) for an explanation of this fact based on the existence of moral hazard and risk of repudiation). One additional piece of often cited evidence against international consumption insurance is the fact that

the portfolio of developed countries is composed in large part of domestic assets (see e.g. Tesar and Werner (1992)). Since there are instances when agents may be better off by changing the composition of their portfolio (see e.g. French and Poterba (1991)), it appears that they fail to reap the benefits of international diversification. While the question of why portfolios are so little diversified is a subject of an intense debate, several authors have provided a simple rationalization of this phenomenon. For example, Stockman and Dellas (1989) show that nondiversification may occur if nontraded goods account for a large fraction of total consumption. Alternatively, Cole and Obstfeld (1991) and Obstfeld (1992a) have suggested that, given existing frictions, the gains from international diversifications may be small. This may occur because there are large informational costs involved in predicting future payoffs of foreign assets and/or because the cyclical properties of national incomes across developed countries are alike (see e.g. Backus and Kehoe (1992)) (see e.g. van Wincoop (1992) for an opposite view regarding the size of the gains).

The purpose of this research is to empirically examine whether the consumption insurance proposition holds at an international level. We begin in section 2 by showing the implications of international risk sharing in a very simple theoretical setup. We then describe the implications of consumption insurance in more complicated models and show that monotonic transformations of aggregate consumption must be highly correlated across countries, even when preferences are time nonseparable, when there are nonseparabilities across goods, when leisure choices are included and when there are nontraded goods. In this section we also derive the testable implications of the theory and compare our testing approach to others which exist in the literature.

The third section of the paper is devoted to empirically check the validity of the theory. With the exceptions of Obstfeld (1989), (1993) and Lewis (1993), the existing work tests risk sharing by computing the correlation between US consumption and consumption of some other country (typically Japan or Canada) with their respective cross country output correlation. Because consumption correlations tend to be lower than output correlations, international consumption insurance appears to be violated, a phenomenon that Backus, Kehoe and Kydland (1993) describe as one of the most important unexplained puzzles in the international real business cycle literature. The reported evidence, however, concerns primarily the US and one may expect closer consumption ties between countries within the EEC community or between countries where labor migrations have created a self–insurance mechanism of transfers to chronically depressed areas. In addition, since no standard errors are typically provided, comparisons of point estimates of the contemporaneous correlations need not provide conclusive evidence on the issue.

In section 3 we study the properties of pairs of aggregate consumptions for a panel of nine fairly homogeneous developed countries over the sample 1970–1990. We test the relevance of the theory of international risk sharing using a GMM technique. Since the theory predicts that aggregate domestic consumption within countries should not covary with country specific shocks once we control for aggregate foreign consumption, natural

orthogonality conditions emerge by interacting domestic variables (including variables proxying for real, fiscal, monetary and demographic factors), with the residuals of the regression between pairs of aggregate consumptions.

Because our simple model disregards issues connected with the allocation of resources over time, we need to abstract from permanent income considerations in conducting tests of international consumption insurance. In other words, in testing the risk sharing proposition, we wish to consider only those fluctuations which are stationary and this brings up the issue of which stationary inducing transformation to employ. Canova (1992) shows that different detrending methods leave cycles of different length in the data. By appropriately selecting the detrending procedure, we can therefore examine whether variables belonging to the information set of agents are correlated with the component of domestic aggregate consumption which has all its fluctuations over short (2–3 years), medium (4–6 years), long (7–10 years) or infinite cycles. A differentiated analysis across frequencies may shed light on two important issues. First, whether measurement errors which are present in consumption data, and may be particularly important at high frequencies, affect the results of the tests. Second, whether consumption insurance primarily covers temporary or more permanent types of disturbances. Marcet and Marimon (1992) have shown that different informational structures lead to optimal risk sharing arrangements that cover different types of cyclical fluctuations. In addition, there is considerable debate in the policy circles regarding the provision of international insurance without permanent income redistribution across countries (see Melitz and Vori (1992)). Our analysis may therefore shed some light on this issue.

The international risk sharing proposition also imposes two additional restrictions: a set of cross–equation constraints across the moments we test and the restriction that consumption correlations across countries must be perfect. Therefore, additional tests of the theory can be undertaken by examining the validity of these restrictions for each pair of countries.

As a byproduct of our empirical analysis, we provide novel estimates and standard errors of the ratios of relative risk aversion coefficients of the representative agent of pairs of countries. These estimates are of independent interest for researchers engaged in calibrating international business cycle models and may provide a rationalization for the low international portfolio diversification observed in developed countries.

Four major results stand out from our empirical analysis. First, aggregate consumption appears to be fully insured against *shocks* to real, fiscal, monetary and demographic variables. This is true regardless of the detrending method we consider. Second, aggregate consumption covaries with some lagged country specific demographic and labor market variables over medium–long cycles. These two results are very robust to the presence of measurement errors, modifications of the empirical specification of the model, changes in the set of instruments employed and alterations of the functional form for the utility function. Third, the other two implications of the risk sharing proposition are, in many instances, rejected regardless of detrending methods we considered. However, one should note that consumption correlations among European countries are larger than

among the countries of the rest of the panel. Fourth, we find very little evidence of risk sharing in the very long run, both in the sense that consumption patterns tend to diverge in the long run and that the long run component of domestic consumption reacts to domestic variables belonging to the information set of agents.

Several important conclusions emerge from our investigation which we discuss in section 4. These conclusions have relevant implications for the design of institutions intended to implement first best allocations, for the opening new financial markets and for questions concerning fiscal independence of national governments in a unified Europe. First, because existing arrangements appear to provide enough insurance against domestic business cycle fluctuations, both the opening of international financial markets or the introduction of new government institutions intended to augment the efficiency of existing markets (see e.g. Persson and Tabellini (1992)) would be largely irrelevant for the welfare of the representative agents of these countries. This result is particularly important for European countries since one of the foreseeable tasks of the central authority of a unified Europe is to provide consumption insurance to member countries (see e.g. Padoa–Schioppa Report (1987)). Second, because consumption insurance is harder to obtain for fluctuations of longer length, international institutions (old and new) should be more concerned with shielding consumers from this type of domestic fluctuations. However, this type of insurance may involve semi–permanent redistribution of income. Therefore to fully evaluate the effects of these programs on the well being of agents, it is seems necessary to examine how the two components of policy interact in the context of a specific model of insurance and redistribution.

2 A Theory of International Risk Sharing

We assume that there are J countries in the world, $j = 1, \ldots, J$ and that there is a representative consumer in each country with preferences defined over an aggregate consumption good. Let $s_{\tau t}$, $\tau = 1, \ldots, S$ be a sequence of events where each event is a collection of states of the world which occurred up to t. $s_{\tau t}$ represents the information set available and it is assumed to be observed by the agents of all countries. We assume that S is finite, that the conditional probability of $s_{\tau t}$ being realized is $\pi(s_{\tau t})$ and that $\sum_{\tau=1}^{S} \pi(s_{\tau t}) = 1$. The lifetime expected utility of the representative consumer in each country is given by:

$$V \equiv \sum_{t=0}^{\infty} \beta^t \sum_{\tau=1}^{S} \pi(s_{\tau t}) U[c_t^j(s_{\tau t}), b_t^j(s_{\tau t})] \tag{1}$$

where $0 < \beta < 1$ is the discount factor common to all j's, $c_t^j(s_{\tau t})$ is the consumption of country j at time t, event τ and $b_t^j(s_{\tau t})$ represents all those factors other than consumption which affect agents' utility. These factors may include leisure choices, government consumption, consumption of nontradables, home production effects (as in e.g. Benhabib, Rogerson and Wright (1992)), etc.

We assume that there is only one aggregate consumption good in the world, that

each country has a stochastic endowment $y_t^j(s_{\tau t})$ of the good at each t and that the world social planner takes $b_t^j(s_{\tau t})$ as beyond her control. The problem she faces is:

$$\max_{c_t^j(s_{\tau t})} \sum_{j=1}^{J} \Psi_j \sum_{t=0}^{\infty} \beta^t \sum_{\tau=1}^{S} \pi(s_{\tau t}) U[c_t^j(s_{\tau t}), b_t^j(s_{\tau t})] \tag{2}$$

subject to $\sum_{j=1}^{J} \chi_j c_t^j(s_{\tau t}) = \sum_{j=1}^{J} \chi_j y_t^j(s_{\tau t})$, $c_t^j(s_{\tau t}) > 0$, $\forall j$, $0 < \Psi_j < 1, \forall j$, $\sum_{j=1}^{J} \chi_j = \sum_{j=1}^{J} \Psi_j = 1$. The optimality condition for this problem can be written as:

$$\Phi_j U_{c_{\tau t}}^j = \mu_t \tag{3}$$

where $U_{c_{\tau t}}^j = \frac{\partial U}{\partial c_t^j(s_{\tau t})}$, $\mu_t = \frac{\beta^t \lambda_t}{\pi(s_{\tau t})}$ where λ_t is the Lagrangean multiplier for the resource constraint and where $\Phi_j = \frac{\Psi_j}{\chi_j}$ is a set of time invariant welfare weights which are related to the initial wealth of each country Ψ_j and to the initial size of the country χ_j. Implicit in the formulation of the social planner problem is therefore the assumption that along the equilibrium path the planner will not reallocate wealth and that the size of the country (in terms of population) does not change with time. Also, the problem as stated, can be interpreted as the second stage of a maximization problem where the social planner decides how to distribute world consumption across countries after she has allocated resources optimally between world consumption and world investments.

Since μ_t is independent of j, the above optimality condition implies that for any j, k:

$$\log U_{c_{\tau t}}^j - \log U_{c_{\tau t}}^k = \xi_{jk} \tag{4}$$

where $\xi_{jk} = \log \Phi_k - \log \Phi_j$. Equation (4) states that a monotonically increasing transformation of the marginal utility of consumption of any two countries j and k must be equalized apart from a scale factor and that this condition must hold for all τ and t. A second implication of the optimality condition (3) is:

$$\log U_{c_{\tau t}}^j - \log U_{c_{\tau t}}^a = A_j \tag{5}$$

where $U_{c_{\tau t}}^a = \frac{1}{J} \sum_{j=1}^{J} U_{c_{\tau t}}^j$, $A_j = \log \Phi_a - \log \Phi_j$, $\Phi_a = \frac{1}{J} \sum_{j=1}^{J} \Phi_j$. Equation (5) states that, apart for a scale factor, a monotonically increasing transformation of the marginal utility of consumption of country j is proportional to an increasing transformation of the marginal utility of average world aggregate consumption and that this must hold for all countries j, all states τ and all periods t.

The first implication of optimal risk sharing arrangement refers to the consumption pattern of any pair of countries, the second one to the consumption pattern of one country relative to the average. These two implications are not distinct (taking the difference between any two indices j and k in (5) we obtain equation (4)) and the first is stronger than the second since there is a loss of information in (5) due to the averaging of consumption across countries. That is to say, if (4) holds for all j, k then (5) must be satisfied, while the reverse is not true. For panels with large cross sections it may be convenient to test (5) as it is less cumbersome to implement. However, for a small

panel of countries one can test (4) directly on pairs of countries. Finally, note that since we have assumed that the welfare weights are time invariant, the country specific effect which appears in the constant of (4) and (5) will drop out of the expression when taking the difference between two adjacent time periods.

The implications of equations (4) and (5) are very strong since they equate the marginal utility of consumption for all j, k and for all τ and t. Given the uneven quality of international consumption data, these two implications may fail to hold in practice even though they are true in theory due to different degrees and different serial correlations of measurement errors present in aggregate national data. It is therefore useful to study weaker conditions based on moments of the distributions of the logarithm of the marginal utility of aggregate consumption which may be more robust to measurement errors. From (4) it is straightforward to derive the following moment implications of the risk sharing proposition:

$$E[\log U^j_{c_{\tau t}}] = E[\log U^k_{c_{\tau t}}] + constant \tag{6}$$

$$\text{var}[\log U^j_{c_{\tau t}}] = \text{var}[\log U^k_{c_{\tau t}}] \tag{7}$$

$$\text{cov}[\log U^j_{c_{\tau t}} \log U^j_{c_{\tau t}}] = \text{cov}[\log U^j_{c_{\tau t}} \log U^k_{c_{\tau t}}] \tag{8}$$

Note that (8), together with the fact that $\text{cov}[\log U^k_{c_{\tau t}} \log U^k_{c_{\tau t-q}}] = \text{cov}[\log U^j_{c_{\tau t}} \log U^j_{c_{\tau t-q}}]$, implies that the logarithms of the marginal utilities of consumption for any pair j, k are perfectly correlated. Note also that (7) and (8) hold regardless of the welfare weights employed in the social planner problem as long as they are constant. Therefore, there are two senses in which this last set of conditions is weaker than (4). First, because they imply that international consumption insurance holds for moments of the data instead of point by point in time and state. Second, because these relationships can be examined even when the welfare weights used by the social planner are unknown. In section 3 we will focus on the moment implications of the theory and we will jointly examine (6)–(8).

So far we have not discussed the decentralization scheme which may support the optimal risk sharing allocations. To achieve the optimal allocation in decentralized economies it is common to assume the existence of equity markets in each country (as in Lucas (1982)) or of complete international contingent claim markets where countries trade claims to each others income (as in Backus, Kehoe and Kydland (1992)). However, even when contingent claims markets are missing, sufficient richness in the set of trading opportunities may partially compensate for the lack of variety in the available securities (see e.g. Duffie and Huang (1985)). At the polar extreme, the complete lack of international financial markets is not sufficient to prevent risk sharing. For example, Persson and Tabellini (1992) describe a cooperative political equilibrium arrangement which implements optimal allocations even when international financial markets are absent. Here we leave the form of the financial instruments available unspecified because the presence of financial markets which are efficiently run is only a sufficient condition for the implementation of the optimal allocations and other arrangements, which may involve, for example, bilateral governments agreements or general cooperation agreements, may do the job.

2.1 Setting Up the Tests

To give empirical content to the theory, we specify a parametric functional form for the instantaneous utility function for each j:

$$U(c_t^j, b_t^j) = \frac{1}{1-\sigma_j}[(c_t^j b_t^j)^{1-\sigma_j} - 1] \quad \text{if } \sigma_j \neq 1 \tag{9}$$

$$U(c_t^j, b_t^j) = \log(c_t^j) + \log(b_t^j) \quad \text{otherwise} \tag{10}$$

where σ_j is the coefficient of relative risk aversion of country j. Although we have adopted a CRRA specification, any member of the a general class of HARA utility functions would serve the purpose for the empirical work (see e.g. Brennan and Solnick (1989)). We have chosen a CRRA utility function because of growth considerations. Devereux, Gregory and Smith (1992), who have studied the implications of the risk sharing proposition in an international model of the business cycle where leisure choices are included, have used a utility function of the form:

$$U[c_t^j(s_{\tau t}), \ell_t^j(s_{\tau t})] = \log(c_t - \gamma n_t) \tag{11}$$

For these preferences, the marginal utility of consumption is not independent of labor supply choices but the income elasticity of leisure is zero. Note however that these preferences are not of the HARA class and are incompatible with balanced growth.

As an alternative one could have simply specified a semi–parametric form for the instantaneous utility function. For example, we could have assumed that c_t and b_t interact multiplicatively but left unspecified the functional form for each of them. While this approach is robust to the misspecifications of the functional form for instantaneous preferences, it seems to provide very little gains over a parametric specification in terms of interpretability of the results.

Given the above functional form for preferences, we let

$$X_{1t} = \log c_t^j - (\frac{\sigma_k}{\sigma_j}) \log c_t^k - \xi_1 \tag{12}$$

$$X_{2t} = (\log c_t^j)^2 - (\frac{\sigma_k}{\sigma_j})^2 (\log c_t^k)^2 \tag{13}$$

$$X_{3t} = (\log c_t^k)^2 (\frac{\sigma_k}{\sigma_j}) \text{corr}(\log c_t^k, \log c_t^j) - (\log c_t^j)(\log c_t^k) \tag{14}$$

where $\text{corr}(\log c_t^k, \log c_t^j) = 1$ by the discussion following (8), ξ_1 is a function of the fixed effect of country j and we let $X_t = [X_{1t}, X_{2t}, X_{3t}]'$. If the theory of international consumption insurance holds X_t should be unpredictable. That is to say, for any variable Z_t belonging to the information set $s_{\tau t}$, $E_t[X_t | Z_t] = 0$. Since we have assumed that $s_{\tau t}$ is common to all countries, Z_t may in principle include variables of all countries in the sample. However, since the theory suggests that the log of the marginal utility of consumption is unpredictable given country specific characteristics we will only use current and lagged variables for country j. Note that variables which affect world wealth

should be excluded from Z_t since they are likely to be uninsurable and will therefore induce correlation between the instruments and X_t. An example may clarify this point. Suppose the J countries are non–oil producing and that Z_t includes oil prices. Then $E_t[X_t|Z_t]$ can not be zero when Z_t changes since total wealth has changed. However, if among the J countries there are some oil producing nations, world wealth may stay constant and oil prices may be a reasonable instrument in testing risk sharing.

In theory X_t may include omitted variables or time varying components due to idiosyncratic preference disturbances. In the empirical implementation of the test, X_t may also include functions of the measurement errors in consumption. If the components of X_t are serially correlated, care must be exercised in selecting the Z_t vector because some instruments may be either correlated with some of the unobservable components or measured themselves with an error which is correlated with measurements errors in X_t and this may invalidate the test. For example, if measurement errors in consumption are highly serially correlated, lagged consumption may not be a valid instrument. Similarly, current national income (or current GDP) may not be a valid instrument since its measurement error may be highly correlated with the measurement error in consumption (see also Cochrane (1991)). One solution to these problems is to take instruments sufficiently lagged in the past so as to decrease the extent of the serial correlation in the measurement errors. This procedure has however the disadvantage of also reducing the power of the test since the correlation of the instruments with the current marginal utility of consumption will be smaller. As an alternative, one could simply eliminate from Z_t those variables which are suspected of being correlated with the measurement error in consumption.

While any variable Z_t belonging to $s_{\tau t}$ is a legitimate instrument to examine the basic orthogonality conditions imposed by the theory, the tests conducted using generic Z_t's may be uninformative since there may be time series aspects of Z_t which need not be insurable. That is, the theory of international risk sharing predicts that the logarithm of the marginal utility of domestic consumption is insensitive to idiosyncratic *shocks* after we account for its dependence on the logarithm of the marginal utility of foreign consumption (see also Mace (1991)). A more powerful test of the theory can therefore be constructed by using as instruments the component of Z_t which is unexplained by past values of Z_t's. To construct the unpredictable component of Z_t we use a VAR methodology and obtain shocks by projecting country specific variables proxying for real, fiscal, monetary and demographic factors on their past values. These innovations are then used as instruments to test the international risk sharing proposition.

As is clear from the definition of the X_t vector, the theory of international risk sharing also imposes restrictions across the three orthogonality conditions we examine. Risk sharing, in fact, implies that the slope coefficient of all three moments (12)–(14) is equal to the ratio of the coefficients of relative risk aversions of the representative agents of the two countries. These restrictions are distinct from the basic orthogonality condition $E[X_t|Z_t] = 0$, are independent of the functional form for utility we use (see (6)–(8)) and therefore provide a further and more stringent test of the theory. To empirically

examine their validity we employ a Wald test using the unrestricted estimates of the slope coefficient of each of the moment conditions.

Finally, risk sharing implies that consumption correlations for any pair of countries should be perfect. We test this third implication using a t–test and the point estimates and the standard deviations obtained from the GMM estimation of the three orthogonality conditions.

2.2 Extensions

The restrictions on international aggregate consumption we derived in the previous subsections are obtained from a very simple model and it is legitimate to ask what kind of restrictions are implied by international risk sharing in more general setups. In this section we briefly show that the basic restrictions of the model are very general and hold, in a somewhat modified form, in richer setups.

We consider first the case where preferences do not display time separability and specialize the instantaneous utility function to be $U(c_t^j, b_t^j) = \log(C_t^j - \alpha C_{t-1}^j) + \log(b_t^j)$, where α can be either positive (habit persistence) or negative (durability) (see Mace (1991)). It is easy to check that (4) holds but now the marginal utility of consumption depends on current and past levels of consumption. Note also that, because versions of the Porteus–Kreps preferences are observationally equivalent to von Neumann–Morgenstern preferences with a particular form of habit persistence (see e.g. Constantinides (1991)), the basic implications of risk sharing are also robust to alterations of the von Neumann–Morgenstern assumption on agents preferences.

The restrictions we derived are robust to modifications of the assumption that there is only one good in the model. For example, c_t^j may be an aggregate of j goods, each of which is produced in only one of the j–countries. As Cole and Obstfeld (1991) suggest, as long as there is complete specialization, i.e. country j receives an endowment of good j only, pairs of aggregate consumption should still satisfy (4). However one should stress that, in this setup, country specific disturbances change the terms of trade and these changes may automatically pool national risk. For example, in Cole and Obstfeld's model a technology shock induces, other things being equal, a negative relationship between terms of trade and output. We can also relax the exchange economy setup, introduce production, assume that production requires capital goods from one or more countries and still produce the result that, given the logarithm of the marginal utility of consumption of country k, no variable other than a constant may be important in predicting country j logarithm of the marginal utility of consumption (see Cole and Obstfeld (1991)).

As already mentioned, our framework is general enough to allow for the presence of variables other than consumption in the utility function of agents. The only crucial restriction that is needed for (4) to be satisfied is that all these additional effects (which are taken to be beyond the social planner's control) enter multiplicatively in the utility func-

tion. For example, if there are tradable and nontradable goods in the economy and the instantaneous utility function is of the form: $U[c_t^j(s_{\tau t}), b_t^j(s_{\tau t})] = \frac{1}{1-\sigma_j}((x_t^j)^{\gamma_j}(nt_t^j)^{1-\gamma_j})^{1-\sigma_j}$ where nt_t^j is the consumption of the nontradable good produced in country j, x_t^j the consumption of tradables in country j and γ_j is the share of tradables in total consumption in country j, the optimality condition is

$$\log(x_t^j) - \frac{\gamma_k(1-\sigma_k)-1}{\gamma_j(1-\sigma_j)-1}\log(x_t^j) = \frac{1}{\gamma_j(1-\sigma_j)-1}[(1-\gamma_j)(1-\sigma_k)\log nt_t^k - (1-\gamma_j)(1-\sigma_j)\log nt_t^j] + \frac{1}{\gamma_j(1-\sigma_j)-1}(\log\Phi_k - \log\Phi_j + \log\gamma_k - \log\gamma_j) \quad (15)$$

From equation (15), it is clear that when the marginal utility of consumption of the tradable good is nonseparable in the consumption of the traded and nontraded goods, fluctuations in the endowment of domestic and foreign nontradables will affect the optimal allocation of the traded good. In particular, the optimal allocation will depend on the elasticity of substitution between traded and nontraded goods (γ_j), the intertemporal elasticity of substitution ($\frac{1}{\sigma_j}$) and the joint stochastic processes for traded and non traded goods (see e.g. Tesar (1993)).

The argument we have just described also holds if, rather than nontradables, we assume that there is an exogenous government consumption expenditure affecting the utility of the agents of their own country. The optimality condition of the social planner problem in this case requires equalization of the marginal utility of private consumption across countries. The logic of this case is identical to the one previously described because private and government consumption expenditure may enter in a nonseparable way in the utility function and government consumption cannot be reallocated across countries.

Another case of interest is one where leisure is a choice variable for the agents (and for the social planner). To make the problem meaningful we also abandon the pure exchange economy setup and assume that y_t^j is produced with domestic labor according to a general decreasing returns to scale production function satisfying standard regularity conditions. Let $y_{jt} = f_t^j(1 - L_{jt})$ where $(1 - L_{jt})$ is the amount of labor of country j used to produce y_{jt} and f_t^j is time varying because of a country specific technology disturbance. We assume an instantaneous utility function of the form $U[c_t^j(s_{\tau t}), b_t^j(s_{\tau t})] = \frac{1}{1-\sigma_j}((c_t^j)^{\gamma_j}(L_t^j)^{1-\gamma_j})^{1-\sigma_j}$ where γ_j now represents the share of consumption in the utility function of country j. Then the optimality conditions of the social planner problem can be summarized as follows:

$$\log c_t^j - \log c_t^k = \log L_t^j - \log L_t^k + \log f_{jt}' - \log f_{kt}' + B_{jk} \quad (16)$$

where f_{jt}' is the marginal product of labor in country j and where $B_{jk} = \log\frac{(1-\gamma_k)}{\gamma_k} - \log\frac{(1-\gamma_j)}{\gamma_j}$. The modified condition for international consumption insurance therefore states that differences in consumption profiles across countries should be matched by

differences in leisure and productivity profiles. No other variable should be important in predicting differences in aggregate consumption across pairs of countries.

One final case which is of interest is when there is some heterogeneity within each country. In this situation one can assume that the social planner maximizes a weighted average of the expected utility of the median voter of each country (as in Persson and Tabellini (1992)). Then the international risk sharing condition (4) still holds but now the expressions involve the marginal utility and the individual characteristics of the median voter in each country.

In all the extensions we have presented here, testing the international consumption insurance proposition with private aggregate consumption data is still valid as long as the instruments are judiciously chosen. In other words, if an econometrician omits variables which are either nonseparable with consumption or enter in the computation of the optimal social planner rule, she may reject the risk sharing proposition even if it holds true. In the case of time nonseparable preferences this would be the case if lagged consumption is an element of the instrument set. When there are tradables and nontradables which are nonseparable in the utility function, instruments like domestic income or domestic production would be significant in predicting country j consumption of tradables because they may be correlated with consumption of nontradables in country j (as noted e.g., by Dellas and Stockman (1989)). Finally, if leisure is a choice variable and one implements the test using data only on aggregate domestic consumption, one may find that variables correlated with the marginal product of labor or with leisure may be significant in explaining differences in private consumption across countries.

One way to check whether this source of misspecification is crucial for our conclusions is to include in Z_t all these variables at a second stage and check whether the basic empirical results are altered. If the results do change, one can respecify the utility function in order to capture these effects. If they are robust, we can conclude (i) that, if they exist, these effects are minor and (ii) that failure of risk sharing to hold are not due to comovements with idiosyncratic fluctuations in these variables.

2.3 A Comparison With the Existing Literature

Although the issue of consumption insurance has a long history (see e.g. Wilson (1968)) and has been dealt with theoretically in many research papers, relatively few formal empirical investigations of the implications of the consumption insurance proposition have been conducted. Townsend (1989), Mace (1991), Cochrane (1991) and Atkeson and Bayoumi (1991) have investigated whether consumers within communities, counties or states respond only to aggregate but not to idiosyncratic shocks. The former two studies use individual income as a proxy to measure idiosyncratic shocks and investigate the consumption insurance proposition using pooled time series cross sectional regressions. The latter two employ a number of individual characteristics to represent idiosyncratic shocks and test the insurance proposition using cross sectional regressions. While Townsend and Atkeson and Bayoumi find some evidence of pooling idiosyncratic

shocks, Mace and Cochrane find that there are also many sources of individual risk (such as long illnesses or long unemployment spells) which can not be insured by individual consumers.

At an international level, and to the best of our knowledge, only Obstfeld (1989), (1993) and Lewis (1993) have formally examined the risk sharing proposition. The results of their investigations are mixed. They detect some evidence of market incompleteness but they also find that risk sharing has increased after 1973. Because our model specification, estimation technique, testing approach, data set and panel employed differ from theirs, our investigation provides an alternative and complementary framework to evaluate the usefulness of the complete market abstraction in characterizing international consumption allocations.

Many authors (see e.g. Backus, Kehoe, Kydland (1992), Devereux, Gregory and Smith (1992) or Baxter and Crucini (1992)) have provided informal evidence suggesting that the cross country consumption correlations are less than perfect and, in general, lower than the cross country correlations of national outputs. This empirical evidence combined with simulation results obtained in models where perfect risk sharing occurs have lead Backus, Kehoe and Kydland (1993) to label the cross country relationship between consumptions and incomes a major puzzle. While this evidence is suggestive of the possible lack of international risk sharing, there are at least three reasons why these results are not necessarily conclusive. First, perfect consumption correlations is only one of three possible implications of risk sharing. Second, no statistical conclusions can be drawn because standard errors for the correlations are not provided. Third, as we have already argued in previous sections, low private cross country consumption correlations do not necessarily imply that international risk sharing is absent, especially when nontraded goods, government expenditure or leisure appear in the utility function of the representative agent.

Some authors have also provided more direct evidence on the lack of risk sharing across developed countries by calculating the welfare costs of closing international financial markets or restricting capital flows and showing that the costs are large (see e.g. Brennan and Solnick (1989)). The conclusions regarding the magnitude of the welfare costs of limiting trade in financial markets is, however, the object of current dispute, see e.g. Cole and Obstfeld (1991), Mendoza (1991), Obstfeld (1992a) and van Wincoop (1992). In addition, recent work by Baxter and Crucini (1992) shows that, within standard versions of real business cycle models, the presence of an integrated market for uncontingent government bonds may be sufficient to provide almost complete insurance from country specific technology shocks. Finally, the extent of the losses due to the lack of private risk sharing or the lack of a complete array of international financial markets may also be upward biased if formal and informal government agreements and the terms of trade changes provide semi–automatic insurance (stabilization) schemes for domestic residents.

Other authors have focused on the capital mobility aspect of the risk sharing proposition (see e.g. the literature following Feldstein and Horioka (1980)) or on its implications

for real interest rate differential (see e.g. Obstfeld (1986)). It is often argued that when international risk sharing occurs, capital mobility should be high, real interest rate differentials should be equalized across countries, portfolios should be diversified and the correlation between domestic savings and domestic investment should be low. While these implications are not necessarily true when some consumption goods are nontraded and, as Backus, Kehoe and Kydland (1992) and Baxter and Crucini (1993) have argued, high saving and investment correlations need not indicate imperfections in capital markets, the existing evidence on these issues is somewhat contradictory (see van Wincoop and Marrinan (1993)). While capital mobility appears to be small, in the sense that current account imbalances are small on average for a cross section of countries, riskless bond markets appear to be substantially integrated and the cross country real interest rate differentials appear to be very small.

3 An Empirical Investigation

3.1 Business Cycle Implications

Before we empirically investigate whether aggregate consumption data conform to the theory of international risk sharing, there is one preliminary step that needs to be addressed.

Restrictions (6) through (8) hold for any frequency of the spectrum as long as aggregate consumption is stationary. Since there are many detrending filters which make consumption stationary, it is worthwhile to ask which procedure is the most appropriate to use in testing the theory. Canova (1992) has demonstrated that different detrending methods leave stationary cycles in the data which have different average length. In particular, that paper shows that a first order differencing (FOD) filter leaves in the data cycles with a periodicity which is less than 3 years, that the Hodrick and Prescott (HP) filter leaves in the data cycles with a periodicity between 4–6 years and that a linear detrending method (LT) leaves very long cycles (7–10 years periodicity) in the data. To illustrate the features of these three detrending procedures, we plot in figure 1 detrended US consumption data obtained with each of the three filters. Hence, by appropriately selecting a particular detrending method, one can investigate the consumption insurance relationship across various frequencies of the spectrum.[1] Such a differentiated analysis may shed light on two issues. First, whether the relationship is stronger when idiosyncratic disturbances have a very temporary or a more permanent nature. Second, whether the large component of measurement error present in consumption data at high

[1]It is useful to stress that an approach that separates the information by frequency shares some similarities with "band spectrum regression" technique of Engle (1974), and with the procedure used in Canova and Dellas (1993), where one computes the average coherence between pairs of series in a selected frequency band and checks whether the value of the statistic over the band can be related to interesting variables.

frequencies is responsible for some of the dramatic results presented in the literature (see e.g. Devereux, Gregory and Smith (1992)).

In our empirical investigation we consider quarterly per–capita total real aggregate consumption data for 9 OECD countries: Australia, Canada, France, Italy, Japan, Switzerland, United Kingdom, USA and Germany. We concentrate on these nine nations because they are sufficiently homogeneous as a group and because they are the only ones with consistent quarterly data for a sufficiently long period. The sources of the data, its definition and its sample availability are described in the appendix. Per–capita total real aggregate consumption may not be the ideal measure of consumption to use in our tests especially if there are persistent cross country differences in the various components of aggregate consumption (such as services, which account for a large portion of nontraded goods) or if the consumption of some countries has more durability aspects than others. As an alternative, one could use consumption data on those non–durable goods which are easily tradable and transportable. There are two drawbacks to this choice, however. First, disaggregated consumption data at a quarterly frequency exist only for a very limited number of countries, so that the power of our cross country comparison would be reduced. Second, to implement the test using non–durable (or tradable) consumption one has to make the assumption that the effect of non–durables (or tradables) is separable from the effect of other components of consumption. If this assumption is not satisfied, the misspecification introduced in the testing procedure may well be larger than the one introduced using total consumption. Finally, one should also remember that various categories of consumption are not necessarily measured in the same way and at the same time in each country. In other words, any available measure of consumption is likely to be unsatisfactory since numerous sources of misspecification may be introduced in the testing procedure. Therefore, one should be very cautious in interpreting the results of the tests.

Tables 1 and 2 present the basic results of our investigation. Table 1 reports the significance level (in percentage) of a χ^2 test for the overidentifying restrictions implied by the risk sharing proposition when we jointly use the three moment conditions (6)–(8) and the instruments are a constant, domestic real variables (lagged outputs), domestic demographic variables (lagged population) and domestic nominal variables (lagged prices). All data is logged before any transformation undertaken. The first panel of the table presents the results for LT detrended data, the second panel for HP detrended data and the third panel for FOD detrended data. In all the tests presented in this table we apply to the instruments the same detrending transformation we employed on the consumption data. That is, if we linearly detrend consumption, we also linearly detrend the instruments to conduct the test. This allows us to concentrate on the comovements between consumption and the instruments at a particular frequency and makes the cross frequency comparisons more appropriate. In conducting the tests we only use lagged values of the instruments (as opposed to current and lagged values) in order to minimize endogeneity problems, which may be present primarily with output (or income) data, and measurement error problems, which may arise with all types of instruments if their

measurement errors are correlated with measurement errors in consumption. Table 2 presents the results of testing the overidentifying restrictions implied by the risk sharing proposition when we still use the three moment conditions (6)–(8) but the instruments are a constant and current and lagged *shocks* to real, demographic and nominal domestic variables. Current shocks are used because they are exogenous to agents' information set and they are unlikely to be correlated with the measurement error in consumption. Shocks are constructed by taking the residuals of a VAR(4) on domestic output, domestic population and domestic prices.

Two main results stand out from these tables. First, aggregate domestic consumption is insensitive to idiosyncratic shocks in domestic instruments, a result which is consistent with international consumption insurance. Second, aggregate domestic consumption seems to covary with some of the domestic instruments but the strength of this association depends on the length of the cycles considered. For example, over short cycles, the risk sharing proposition is never rejected while for longer cycles, there are some comovements between domestic consumption and some of the instruments.

In order to understand why the result differs in the two tables, it is important to identify which instrument covaries with domestic consumption. It turns out that over medium–long cycles, aggregate consumption covaries primarily with demographic variables while it is essentially unaffected by movements in the other instruments. This should not come as a surprise since the tests are performed on per–capita consumption data and population, which is highly serially correlated within each country and does not move in a similar way over the medium long run across the nine countries. In addition, if we note that innovations in demographic variables are small over the sample under consideration, we find that the results of the two tables are very consistent.

There are at least four explanations for why the strength of the association between consumption and the instruments varies with the detrending method. First, it may be the case that aggregate consumption is insensitive to fluctuations in domestic variables when these movements are very temporary in nature. Marcet and Marimon (1992) showed that in an economy with participation constraints this will be the case: agents will insure only short run fluctuations so that domestic consumption will track permanent income in the medium–long run. Alternatively, it may be the case that fluctuations which last longer may be harder to insure against because they may signal the presence of structural imbalances in the economy. Second, it is well known that aggregate consumption data tend to be well approximated by random walks. That is, the cyclical component of aggregate consumption which is extracted using a FOD filter is, for most practical purposes, unpredictable given available information. In this case the tests performed using FOD detrended data may be very weak, since no variable, either domestic or foreign, is useful in predicting consumption movements over short cycles. We do not believe this explanation entirely accounts for the results for two reasons. First, when we use FOD detrended data, the estimated coefficient on the foreign aggregate consumption is significant in almost half of the cases, therefore contradicting the idea that the log FOD of consumption is unpredictable. Second, when we use HP detrended data,

foreign aggregate consumption always has a significant coefficient but we still reject the risk sharing proposition in about one–third of the cases. In other words, there appears to be little relationship between the significance of the coefficient on foreign consumption and the rejection of the risk sharing proposition so that the above explanation is, at best, suggestive. The differences we observed can also be explained with the fact that measurement errors may be less correlated across variables at high frequencies than at low frequencies. This could be the case, for example, if it is easier to measure growth rates than levels. In this situation the pattern of covariations at longer cycles can be attributed to cumulative errors which are made in measuring variables. Although in principle this explanation has some appeal, it is hard to quantify how important this type of problem is in the present context. A final possible explanation for the pattern of results has to do with the type of consumption data we are using. Because disaggregated data for many countries do not exist, we are forced to test international risk sharing using total aggregate consumption. However, this may create problems in interpreting the results because consumer durables are more likely to respond to more permanent fluctuations in the instruments. In other words, if important components of total aggregate consumption have some durability aspect, it may turn out that aggregate consumption is acyclical over very short cycles but that it reacts to the state of the economy when the length of the fluctuations increases. While this fact constitutes a matter of concern, it should be noted that it is not fully satisfactory as an explanation because the pattern of cross–frequency results we found is robust to the choice of the pair of countries despite the fact that the durable component of total consumption is not very homogeneous across countries.

To check the robustness of our conclusions, we conducted several sensitivity exercises. We have two goals in mind: first, we are interested in knowing whether the pattern of results depends on the choice of instruments or on some of the auxiliary assumptions we have made in performing the tests. Within this class of exercises we checked whether the presence of serially correlated measurement errors is responsible for the pattern of rejections present in table 1 by changing the timing of the instruments and performing the tests using current and lagged instruments. To quantify the importance of measurement errors for the tests we added current income and lagged consumption to the instruments. To examine whether the particular choice of variables is responsible for the results we also modified our proxies for real, demographic and monetary instruments and included next exports as an instrument. To check whether the presence of a country specific effect mattered in the tests we omitted a constant in testing (6). Finally, to examine whether risk sharing has increased after 1973, we rerun all the tests using data from 1974 to 1990.

Second, we study whether a misspecification of the empirical relationship is responsible for some of the results we obtain in table 1. In section 2.3 we showed that when government expenditure or leisure enters the utility function, the risk sharing proposition implies that consumption across countries should be equalized after taking into account terms due to government expenditure, hours or productivity of labor. It is possible that the instruments we used are correlated either with government expenditure or hours,

so the rejections we observe may be the result of misspecifications. In addition, it may be that the CRRA specification for the utility function of the representative agent is incorrect and a (locally) quadratic utility function is more appropriate. In that case international consumption risk sharing implies that the cyclical component of consumption (as opposed to the cyclical component of the logarithm of consumption) should be proportional across countries. To examine whether these forms of misspecification are present, we added lagged government expenditure and lagged employment to the set of instruments and also ran the tests on the level of the cyclical components of consumption.

The results obtained in tables 1 and 2 appear to be surprisingly robust and none of the conclusions is affected by any of our modifications. Particularly important is the robustness of the results to the elimination of an exponential trend in the data (which contrasts with the results of Mace (1991) at a micro level) and to the exclusion of employment and government expenditure from the instruments, which suggests that a misspecification of the arguments of the utility function is not the reason for the failure of the risk sharing proposition over long cycles.

One other important implication of the above results is that government consumption expenditure is essentially acyclical in all the countries of the panel. This may signal the lack of direct programs aimed at insuring private consumption through automatic movements in government expenditure and it is in contrast with some of the results of Christiano and Eichenbaum (1992) who suggest that government expenditure is an important determinant of cyclical fluctuations in aggregate consumption in the US.

As emphasized in section 2, apart from orthogonality condition, the international risk sharing proposition also implies a set of cross equation restrictions on the estimates of the slopes of (6)–(8). These restrictions are independent of the exact parametrization of the utility function and can be used to form an alternative and possibly stronger test of the theory. To check whether the estimates of the slopes obtained from three different moment conditions are the same we use a Wald test. The results are very strong. Regardless of the instruments used or of the frequency considered, the cross equation restrictions are always soundly rejected. Hence, although aggregate consumption does not covary with idiosyncratic domestic shocks, the stronger implication that (6)–(8) provides the same information does not hold true.

Despite the rejection of the cross–equation restrictions, our estimates of the slope coefficients in (6)–(8) are of independent interest for several reasons. First, they may be useful to researchers engaged in calibrating international business cycle models since they provide a measure of how risk averse a fictitious representative agent of each country is relative to the representative agent of a base country. Second, they may provide some rationale for the apparent lack of international diversification in the portfolio of different countries (see e.g. French and Poterba (1991) and Tesar and Werner (1992)).

To give a direct meaning to the slope estimates, and because more is known about the estimate of this parameter for US consumers, we use the relative risk aversion of the representative consumer in the US as a numeraire. Table 3 presents point estimates of the

relative risk aversion coefficients obtained with (7) and their standard errors. We chose to present only the estimates obtained with (7) because they are more reasonable given that in the estimation they are constrained to be positive. We confine point estimates of the relative risk aversion obtained with the other two conditions to an appendix available on request.

It is immediate to note that the sets of estimates obtained with different instruments are neither statistically nor economically different. In general, it appears that the representative agent of half of the countries are less risk averse than the representative agent in the US. Since the ratio of relative risk aversion coefficients across countries is one of the deep parameters of the model, one would also hope that estimates of these ratios do not differ too much across frequencies. We therefore conducted a formal Wald test to check whether this is the case. The evidence on this issue is mixed. Except for the case of France, Switzerland and the UK, estimates of the relative risk aversion coefficients obtained with LT and FOD are not significantly different. On the other hand, except for the case of UK and Switzerland, estimates of the relative risk aversion coefficients obtained with HP and FOD are significantly different. Finally, in all cases, estimates of the relative risk aversion coefficients obtained with HP and LT are significantly different. One should however note that even though the test rejects the hypothesis that the three sets of coefficients are the same, except in the case of Switzerland, the range of estimates obtained over different frequencies is economically small.

Obstfeld (1989) has also provided estimates of the risk aversion parameter for the US, Germany and Japan. The point estimates he reports are similar to ours but he finds that over the sample 1961–1985 they are significantly different from each other, while over two subsamples (1961–1972, 1973–1985) they are not. Apart from the different sample used, the major difference between the two sets of results is due to the size of the standard errors. Obstfeld's estimated standard errors are in fact 100 times larger than ours. This indicates that the instruments he chose may have very little information regarding the risk aversion characteristics of each country.[2]

Finally, implicit in the three moment conditions we tested is the condition that consumption correlations are perfect. While the rejection of the cross equation restrictions contained in (6)–(8) suggests that consumption correlations are far from perfect, it is useful to examine the size of the estimated correlations which are implicit in the estimated slope coefficients of (8) for two reasons. First, to formally test whether estimates of the correlation coefficient are different than 1. Second, because their magnitude may explain some of the differences between the results presented here and existing work. Table 4 presents the estimated correlation coefficients and the estimated standard errors when the instruments used are detrended domestic variables. Estimates obtained when the instruments used are shocks to domestic variables are very similar and not reported here. Three major results stand out from the table. First, consumption correlations are

[2]To confirm our intuition we also conducted tests where we constrain the risk aversion coefficient to be the same in each country. As expected, this increases the rejection rate for all specifications we tried indicating that the restriction is not satisfied in this data set.

significantly different from one in almost all cases for all types of cycles. This therefore confirms the informal analysis of Backus, Kehoe and Kydland (1992) and Devereux, Gregory and Smith (1992). Second, consumption correlations are stronger and in some cases larger than output correlations for EEC countries, a result which suggest that countries with closer economic ties may also have more efficient consumption risk sharing mechanism. Third, consumption correlations obtained with the UK as one of the country pairs are negative when LT detrended data are used, a result which suggests the presence of high durability content in UK consumption data (see also Blackburn and Ravn (1992)).

3.2 Long Run Implications

The risk sharing proposition is an appropriate idealization to describe a mechanism which insures cross–country aggregate consumption from domestic *cyclical* fluctuations. However, in the formulation we have described in section 2, there is nothing that prevents us from considering other types of fluctuations. In other words, it is also possible to investigate implication (6) of the risk sharing proposition even when domestic cycles have infinite length. In this case the consumption insurance relationship requires (i) that the aggregate consumption profile of pairs of countries should move together in the long run and (ii) that the error in predicting the logarithm of the marginal utility of consumption of country j, given the logarithm of the marginal utility of consumption of country k, should be unpredictable using long run characteristics of country j. Intuitively, the two conditions imply that if international consumption insurance is in effect, the consumption profile across countries cannot permanently diverge. Obviously, as has been pointed out in the literature (see Melitz and Vori (1992)), if this type of insurance is in place, it involves an almost permanent system of income transfers across countries.

An analysis of the long run implications of international risk sharing may also shed some light on a related issue which is of interest to macroeconomists. Following the permanent income hypothesis tradition, it is typical to identify the permanent component of income with the long run behavior of consumption (see e.g. Quah (1990)). Our cross country analysis may therefore shed light on the relationship among permanent incomes across countries and on the issue of convergence of national incomes, a topic which has received substantial attention in the current growth literature (see e.g. Barro and Sala–i–Martin (1992)).

We start by examining a weak long run implication of the risk sharing proposition: that the error in predicting the logarithm of the marginal utility of consumption of country j, given the logarithm of the marginal utility of consumption of country k, is a stationary process. This implication is weak because the error may be stationary but still be predictable given country specific characteristics. The essence of this implication can be easily tested using cointegration tools. We conduct tests for integration and cointegration in aggregate consumption data using Phillips' Z–test[3] and we find that we

[3]For sensitivity we also conducted tests using the Dickey and Fuller (1981) and Stock and Watson (1989) tests. The three procedures produce the same answer for all series except for French private

cannot reject the hypothesis that aggregate consumption is an I(1) variable, except for Switzerland. For each pair of countries for which consumption is an I(1) variable, we then run a regression like $\log c_{jt} = a + b * \log c_{kt} + e_{jt}$ and check if the residuals contain a unit root.[4] Table 5, panel A presents the results of these tests. In the table we report the least favorable outcome, that is we report the largest (in absolute value) t–statistics over possible lags augmentation, and the lag value at which it occurs. The results show that the e_{jt}'s appear to contain a unit root in many cases (exceptions are the pairs Australia–Japan, Japan–France, Australia–France in addition to some borderline cases). Hence, consumption patterns tend to permanently diverge and even this weaker version of the risk sharing proposition is generally unsupported in the long run.

For those countries for which consumption appears to be stationary (and for borderline cases) we next examine the second implication of the proposition. Table 5, panel B reports the results. We find that risk sharing is rejected in the long run in half of these cases and that domestic consumption covaries in the long run with all the domestic variables we consider. Two interesting conclusions can be derived by combining the results of these two tests presented in table 5. First, there are some cross frequency effects with the stationary component of consumption reacting to long term movements in the economy. Second, permanent income across countries show little convergence, a result which appears to contradict the long run implications of the basic neoclassical growth model.

4 Conclusions

In this paper we have analyzed some international aspects of the risk sharing proposition. We showed that risk sharing is a property of optimal international consumption allocations in a wide variety of theoretical setups and we empirically examined whether the theoretical relationship is supported by the data. We find that aggregate domestic consumption seems to be well insured against domestic shocks which are cyclical in nature but that it tends to covary with demographic and labor market variables, primarily over long cycles. We also show that although the orthogonality conditions imposed by the theory cannot be rejected when we consider domestic shocks, we do strongly reject both the cross equation restrictions and the implication that consumption correlations are perfect. We also show that very little risk sharing appears in the long run both in terms of long run comovements of consumptions across countries and of covariations with domestic instruments. As a by–product of the analysis, we have also provided estimates

consumption where the Stock and Watson test rejects the null of a unit root. We also experimented with different lags augmentation and excluding the linear trend from the regression, without important changes in the results. Tables with the results are available on request from the authors.

[4]Because of the poor quality of consumption data and the relative short sample available we also considered an approach which imposes that pairs of consumption series are cointegrated with cointegrating vectors [1, -1] and examine whether or not the residuals display a unit root. Although some of the results change, the basic message of the exercise is maintained.

of the ratio of relative risk aversion coefficients for pairs of countries in the sample, which can be used by researchers interested in calibrating international business cycles models.

Our results are, overall, in agreement with those of Obstfeld (1989), (1993), Atkeson and Bayoumi (1991) and Lewis (1993). However, contrary to them we reject the idea that domestic consumption reacts to several sources of idiosyncratic shocks and we are able to distinguish which of the implications of risk sharing with complete financial markets are at odds with the data. Our conclusions are also in line with those of Backus, Kydland and Kehoe (1992), Devereux, Gregory and Smith (1992) or Baxter and Crucini (1992). We do find that consumption correlations are both statistically and economically different from one but we also find that among European countries the correlations are higher than elsewhere and, in some cases, higher than output correlations, suggesting that some risk sharing is actually taking place.

Although the analysis we have conducted does not specify the market structure which can support the optimal allocations, there are at least two implications of our results which can be useful in designing mechanisms intended to implement first best allocations. We have shown that, whatever they are, existing market structures shield domestic consumption from idiosyncratic shocks sufficiently well (recall that government consumption expenditure does not play a major insurance role). This result seems to deny the need of further government intervention both in terms of providing automatic stabilizers in the economy and new institutions which may help the economies to achieve Pareto optimal allocations. However this statement needs two qualifications. First, the countries in the panel we examined are among the most industrialized of the world and for some of them temporary labor migrations or semi–permanent remittance programs from emigrants may have created an insurance program which need not to be present for other OECD countries or, worse, for LCD countries. Second, because domestic aggregate consumption covaries with some domestic variables over cycles of longer length, there is some room to improve the quality of consumption allocations by designing institutions or opening markets which insure agents against this type of fluctuations. However, insuring fluctuations of longer length may have a marked redistributive effects.

A second issue which is of interest concerns the need of integrating domestic financial markets into a world market. Our results suggest that the opening of foreign financial markets to domestic consumers will be somewhat redundant in terms of welfare as the existing structures offer sufficient insurance against the most interesting sources of domestic fluctuations. Hence, the prospective integration of European financial markets and the current globalization of security markets are unlikely to bring substantial changes in the features of domestic consumption fluctuations.

Data Appendix

The data we use is all taken from Datastream. Consumption measures aggregate private consumption expenditure on nondurables, durables and services. It is transformed into a per–capita series by dividing the original series by population. Because data on population is annual, quarterly data are obtained by taking the predicted values of an AR(3) regression fitted to a dummy quarterly series, constructed assigning the annual value to each of the four quarters. Government data measures current government expenditure except in the case of Australia where also gross government fixed investment is included. Income data measures gross domestic product (GDP) except for Japan, US and West Germany where it measures gross national product (GNP). All data is in real terms. The base year however, differs across countries. For Australia, Italy, Japan, UK, and West Germany the base is 1985, for France and Switzerland the base is 1980, for Canada the base is 1986 and for the US the base is 1987. All variables are measured in annual rates. Employment data is not completely compatible since it measures different aggregates in different countries. The series used measure total employment in Canada, Italy, Japan, Switzerland, UK, USA, West Germany, employment on the payroll in Australia and civilian employment in France. Because some of these series are non–seasonally adjusted, we deseasonalize them by using an exponential smoothing procedure. UK employment data are the same as those employed by Blackburn and Ravn (1992). Finally, the price data measures the implicit price deflator for GNP (or GDP).

The sample we have available covers the period 1960,1–1991,4 for Australia, Canada, United Kingdom, USA and Germany; the period 1965,1–1991,4 for Japan; the period 1967,1–1991,4 for Switzerland; and the period 1970,1–1991,4 for France and Italy. The tests are constructed using the shortest of the two data samples for each pair.

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Table 1
P–values for the J–test (in percentages): Detrended Instruments

	AUS	CAN	FRA	ITA	JAP	SWI	UK	US	WG
Linearly Detrended data									
AUS		0.0143	0.0396	0.1008	0.0791	0.9778	0.0003	0.6023	6.3237
CAN			1.9655	0.0730	0.0131	2.9627	0.0001	0.0000	0.0079
FRA				1.7262	1.9468	6.2842	0.1165	0.5232	0.7088
ITA					0.2206	0.4928	0.8878	2.9091	0.2089
JAP						0.9912	0.0000	0.0161	0.9182
SWI							1.4268	7.5635	1.9050
UK								0.0041	0.0205
US									0.0090
HP Detrended data									
	AUS	CAN	FRA	ITA	JAP	SWI	UK	US	WG
AUS		0.8781	1.1076	12.965	10.040	25.763	7.2284	1.3623	14.541
CAN			15.918	3.3472	7.6228	4.2277	2.5303	0.8536	0.3313
FRA				8.8903	10.872	5.5677	7.3689	15.560	18.961
ITA					12.452	7.4112	3.4685	6.1406	74.793
JAP						16.439	5.1440	1.5182	40.020
SWI							27.766	14.668	14.614
UK								0.1582	4.2776
US									2.4461
FOD Detrended data									
	AUS	CAN	FRA	ITA	JAP	SWI	UK	US	WG
AUS		27.133	24.361	17.586	10.810	34.740	6.0590	51.467	15.893
CAN			74.914	48.487	48.610	15.277	34.339	92.790	18.617
FRA				15.146	49.769	38.609	62.368	43.603	59.375
ITA					18.040	12.192	6.2284	28.422	51.058
JAP						42.652	41.906	49.538	3.8762
SWI							26.881	54.242	39.708
UK								79.449	52.388
US									20.179

Notes: The instruments used to construct the J–tests are a constant domestic output, domestic prices and domestic population lagged one and two periods. The p–values refer to a χ^2 test with 17 degrees of freedom.

Table 2
P–values for the J–test (in percentages): Shock Instruments

	AUS	CAN	FRA	ITA	JAP	SWI	UK	US	WG
Linearly Detrended data									
AUS		81.9882	46.2925	97.6432	73.4789	65.6436	67.3339	65.1412	79.8557
CAN			65.6527	59.7208	42.5670	68.9452	21.3315	26.3213	42.7275
FRA				66.5922	44.5308	34.3820	24.6531	55.7311	31.3766
ITA					39.1124	53.3856	60.1932	37.9021	22.1858
JAP						74.5787	19.1094	93.1501	98.0725
SWI							25.9047	51.7344	10.1648
UK								2.9673	5.8525
US									3.0361
HP Detrended data									
	AUS	CAN	FRA	ITA	JAP	SWI	UK	US	WG
AUS		26.8775	44.3816	81.6847	2.9654	45.9652	29.9638	33.5965	20.1338
CAN			62.5038	75.9179	71.3667	65.4429	45.6255	50.0499	22.6597
FRA				58.9114	60.5837	32.4957	93.5456	56.5847	32.8338
ITA					60.0372	17.886	45.1379	33.8598	70.8611
JAP						52.9332	63.5244	15.1089	76.8317
SWI							54.5571	65.6172	76.2277
UK								50.3640	15.0141
US									10.7206
FOD Detrended data									
	AUS	CAN	FRA	ITA	JAP	SWI	UK	US	WG
AUS		67.8751	43.2496	52.1166	25.2209	68.4601	95.7350	83.7958	99.0131
CAN			7.4336	34.0989	79.2133	31.9034	56.5786	53.2041	42.6540
FRA				73.5163	11.8144	16.1363	54.6077	30.4766	55.0956
ITA					4.0657	18.2037	69.9286	66.8692	12.6185
JAP						60.3460	89.4216	67.3000	46.6399
SWI							21.0858	21.2397	18.6343
UK								46.1325	28.9546
US									86.2626

Notes: The instruments used to construct J–tests are a constant, the current residuals and one period lagged residuals of a VAR(4) on domestic output, domestic prices and domestic population. The p–values refer to a χ^2 test with 17 degrees of freedom.

Table 3
Estimates of the Coefficients of Relative Risk Aversion

	Detrended Instruments			Shock Instruments		
Country	LT	HP	FOD	LT	HP	FOD
AUS	1.006	1.282	0.819	0.947	1.486	0.859
	(0.032)	(0.049)	(0.087)	(0.052)	(0.062)	(0.104)
CAN	0.739	0.857	0.786	0.715	1.049	0.777
	(0.036)	(0.059)	(0.082)	(0.075)	(0.087)	(0.094)
FRA	1.940	1.621	1.170	1.808	1.677	1.138
	(0.027)	(0.016)	(0.046)	(0.040)	(0.030)	(0.068)
ITA	1.912	1.179	1.632	1.305	1.117	1.526
	(0.056)	(0.042)	(0.037)	(0.090)	(0.063)	(0.067)
JAP	0.472	1.020	0.695	0.546	1.045	0.678
	(0.054)	(0.044)	(0.073)	(0.117)	(0.055)	(0.120)
SWI	2.581	1.492	1.132	2.879	1.521	1.303
	(0.017)	(0.039)	(0.039)	(0.035)	(0.071)	(0.072)
UK	0.940	0.689	0.660	0.940	0.689	0.625
	(0.038)	(0.073)	(0.089)	(0.094)	(0.106)	(0.114)
GER	0.571	0.830	0.509	0.571	0.830	0.618
	(0.022)	(0.042)	(0.049)	(0.034)	(0.058)	(0.048)

Notes: The estimates are obtained from the variance restriction (13). Standard errors of the estimates are in parenthesis.

Table 4: Estimated Consumption Correlations

	AUS	CAN	FRA	ITA	JAP	SWI	UK	US	WG
Linearly Detrended data									
AUS		0.767	0.387	0.246	0.809	0.154	-0.514	0.726	0.783
		(0.027)	(0.042)	(0.047)	(0.014)	(0.063)	(0.041)	(0.038)	(0.223)
CAN			0.638	0.393	0.629	0.560	-0.226	0.624	0.809
			(0.044)	(0.047)	(0.035)	(0.056)	(0.071)	(0.032)	(0.020)
FRA				0.671	0.409	0.472	-0.184	0.292	0.586
				(0.042)	(0.054)	(0.061)	(0.061)	(0.055)	(0.057)
ITA					0.194	0.449	-0.112	-0.103	0.274
					(0.023)	(0.050)	(0.041)	(0.057)	(0.041)
JAP						0.367	-0.225	0.443	0.826
						(0.066)	(0.031)	(0.036)	(0.015)
SWI							0.232	0.426	0.383
							(0.064)	(0.052)	(0.053)
UK								-0.113	-0.331
								(0.059)	(0.044)
US									0.551
									(0.035)
HP Filtered data									
AUS		0.217	0.237	0.250	0.232	0.024	0.115	0.047	-0.194
		(0.065)	(0.083)	(0.087)	(0.072)	(0.083)	(0.061)	(0.067)	(0.068)
CAN			0.401	0.244	0.039	0.455	0.392	0.607	0.060
			(0.069)	(0.051)	(0.058)	(0.059)	(0.049)	(0.041)	(0.071)
FRA				0.172	0.428	0.439	0.508	0.580	0.258
				(0.075)	(0.061)	(0.091)	(0.054)	(0.090)	(0.097)
ITA					0.317	0.468	0.395	0.060	0.205
					(0.068)	(0.058)	(0.056)	(0.080)	(0.096)
JAP						0.287	0.552	0.408	0.229
						(0.080)	(0.065)	(0.065)	(0.080)
SWI							0.429	0.451	0.511
							(0.053)	(0.059)	(0.061)
UK								0.366	0.125
								(0.058)	(0.080)
US									0.340
									(0.074)
FOD Filtered data									
AUS		0.419	0.403	0.070	0.251	0.269	0.046	0.313	0.070
		(0.068)	(0.068)	(0.097)	(0.073)	(0.093)	(0.078)	(0.066)	(0.066)
CAN			0.297	0.133	0.071	0.344	0.089	0.531	0.133
			(0.068)	(0.112)	(0.068)	(0.103)	(0.080)	(0.059)	(0.066)
FRA				0.065	0.299	0.355	0.169	0.162	0.305
				(0.071)	(0.068)	(0.073)	(0.092)	(0.057)	(0.064)
ITA					0.090	0.355	0.094	-0.038	0.048
					(0.094)	(0.118)	(0.078)	(0.101)	(0.098)
JAP						0.035	0.077	0.267	0.091
						(0.068)	(0.048)	(0.046)	(0.053)
SWI							0.154	0.172	0.122
							(0.061)	(0.084)	(0.053)
UK								0.189	0.149
								(0.057)	(0.068)
US									0.233
									(0.077)

Notes: Standard errors are in parenthesis.

Table 5.A
t–values of Philips Z–test

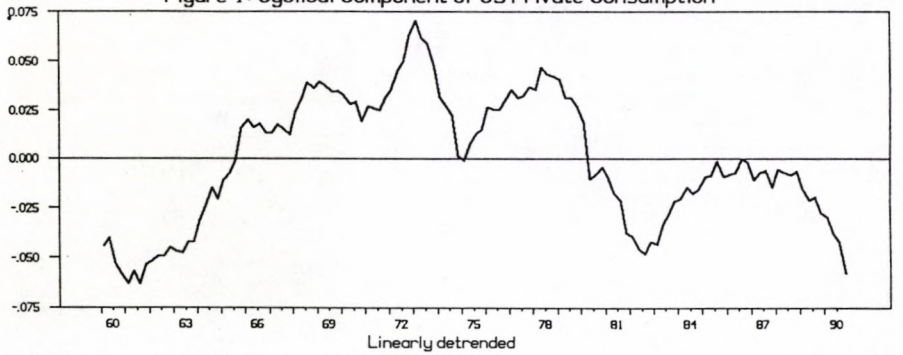
	AUS	CAN	FRA	ITA	JAP	UK	US	WG
AUS		-2.37	-3.88	-3.31	-4.40	-1.69	-2.88	-3.39
		(10)	(6)	(6)	(6)	(10)	(8)	(8)
CAN			-1.89	-2.15	-2.37	-1.37	-2.16	-3.30
			(10)	(10)	(10)	(10)	(10)	(10)
FRA				-3.39	-3.66	-1.93	-2.32	-3.46
				(6)	(10)	(10)	(10)	(10)
ITA					-2.98	-1.63	-2.02	-3.19
					(6)	(10)	(10)	(10)
JAP						-2.97	-2.29	-3.07
						(2)	(10)	(10)
UK							-1.59	-1.65
							(10)	(10)
US								-2.23
								(10)

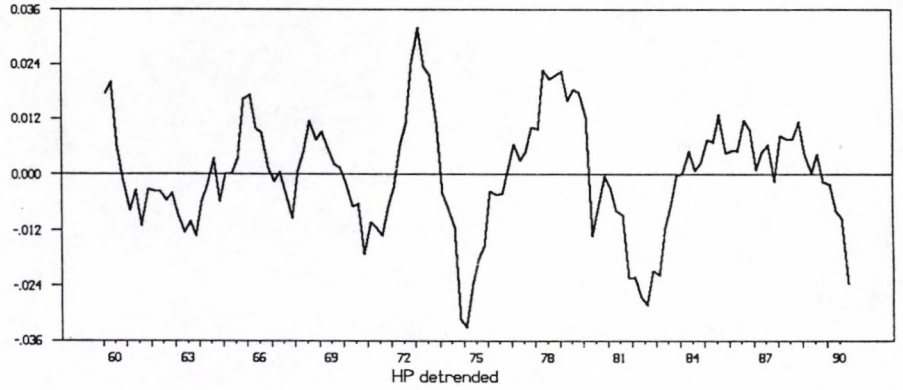
Table 5.B
P–values for the J–test (in percentages): Long Run Cycles

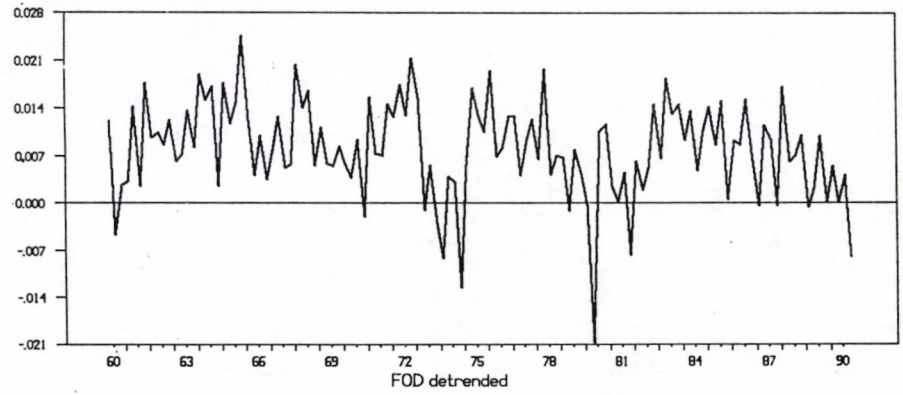
	AUS	CAN	FRA	ITA	JAP	UK	US	WG
AUS			0.2636	0.0000	0.1043			0.0123
CAN								0.0000
FRA				0.2837	0.1265			0.2567
ITA								0.0000
JAP								0.0000
SWI								0.9871

Notes: The instruments used to construct J–tests are a constant, domestic output, domestic prices and domestic population lagged one and two periods. The data is logged before the tests are conducted. Philips' Z–test checks if the residuals of a cointegrating regression between pairs of consumptions are stationary. We report the worst possible case against the risk sharing proposition over possible lag augmentations of the test. The number in parenthesis refers to the lag augmentation used (maximum value is 10). The test is run with no deterministic variables included. Empty cells in panel B indicate that no test is undertaken for that pair.

Figure 1: Cyclical Component of US Private Consumption







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