The Home Bias and Capital Income Flows between Countries and Regions

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This version: June 2008

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Abstract

We suggest using the information implicit in the relative levels of consumption and output to measure long-term international risk sharing. We find that nowadays between 30 and 40 percent of long-term idiosyncratic consumption risk get shared between industrialised countries – as compared to less than 10 percent before 1990. This is a dramatic increase in international risk sharing, in particular if one takes regional evidence from the U.S. as a benchmark, where we find that about 50% of the long-term idiosyncratic risk of the representative state is shared across the Union. The paper identifies a huge increase in international capital income flows as the main driver of this increase in international risk sharing and shows that this rise is linked to the explosion of international cross-holdings of financial assets. While capital income flows remain relatively limited as a channel of risk sharing at business cycle frequencies, our focus on the lower frequency of the data allows us to demonstrate that better international portfolio diversification has indeed led to a considerable increase in capital income flows at medium and long horizons.

Keywords: Consumption Risk Sharing, International and regional business cycles, Capital flows, Home Bias, Non-stationary panel data

JEL classification: E21, F36, F4

1 Introduction

Financial market integration should lead to better international consumption risk sharing. Most of the extant literature has, however, found it difficult to document a consistent effect of financial globalization on consumption. In this paper, we suggest to focus on the long-term comovement of consumption and output in identifiyng this effect. We first show that long-term consumption risk sharing among OECD countries has indeed increased considerably during the recent globalization period, i.e. in particular after 1990. Towards the end of our sample period, which ranges from 1960 to 2004 between 30 and 40 percent of long-term idiosyncratic consumption risk get shared at an international level. Secondly, the paper identifies the channels through which improvements in international risk sharing have come about. We find that considerable increases in medium-term international capital income flows account for the increase in consumption risk sharing and that these improvements can eventually be linked to the dramatic increase in international cross-holdings of financial assets that has been documented in the recent literature.

Our empirical analysis builds on a key prediction from economic theory: standard models with complete financial markets such as the one first studied by Backus, Kehoe and Kydland (1992) imply that fluctuations in relative (i.e. idiosyncratic) marginal utility growth should be independent of idiosyncratic risk. To the extent that marginal utility can be captured by consumption, the coefficient of a regression of relative consumption growth on indicators of idisyncratic risk, such as relative output growth rates, should therefore be close to zero.[1] This coefficient – typically between zero and

[1] Similar regressions were first proposed by Mace (1991), Cochrane (1991) and Townsend (1994) as tests of the null of market completeness. In the macroeconomic literature, they were popularized by Asdrubali, Sørensen and Yosha (1996), Crucini (1999)

one in the data – has subsequently been interpreted as the fraction of idiosyncratic risk that remains unshared, e.g. because financial markets are incomplete: if the coefficient is unity, no risk is shared, if it is zero, all risk is shared. Generally, such risk sharing regressions have been estimated with data that has been rendered stationary through first differencing. An important novelty of our approach is to use the information implicit in the levels of relative consumption and output. By running (panel) risk sharing regressions in relative log levels rather than in relative growth rates, we document longer-term trends in consumption risk sharing that most earlier specifications, based mainly on annual first differences of the data, have not been able to pick up. Based on our levels specification, we show not only that consumption risk sharing has indeed increased considerably since 1990, but that it is actually associated with a considerable growth in international capital income flows. Furthermore, we provide evidence that both phenomena are significantly linked to the growth in OECD countries' international investment positions.

Our results add a novel perspective to what has come to be considered as a key stylized fact on international risk sharing: from the seminal papers by Asdrubali, Sørensen and Yosha (1996) and Sørensen and Yosha (1998) we know that capital income flows derived from cross-holdings of claims to productive capital are a much more important channel of risk sharing among regions or federal states than among countries. Indeed, following Sørensen and Yosha (1998), a number of authors (Lane (2001) and Becker and Hoffmann (2006)) have shown that the fact that this channel is virtually absent in international data can account for virtually all of the lack of consumption

and others. For an excellent survey on home bias and consumption risk sharing see Lewis (1999).

risk sharing at the international level.[2]

The reason why our results differ from most of the earlier literature is that our level risk sharing regressions emphasize the medium and longer term. There is a range of theoretical and empirical justifications for this focus. First, the welfare gains from better risk sharing are likely to be substantial if idiosyncratic shocks are persistent, which is likely to be the case in the data (see also Athanasoulis and van Wincoop (2000) and Becker and Hoffmann (2006)). To the extent that financial globalization lowers the cost of international portfolio diversification more or less uniformly across asset classes, we should expect that improvements in risk sharing would first show up at longer horizons. This is what we see in the data.

Secondly, we would expect the focus on the longer term to allow us to separate the effect on the various risk sharing *channels* more sharply. At business cycle frequencies and as long as macroeconomic fluctuations are transitory it will not make a big difference for consumption allocations whether countries smooth their consumption through savings and dissavings or through capital income flows derived e.g. from equity (Baxter and Crucini (1995)). However, if shocks are sufficiently persistent or even permanent, intertemporal consumption smoothing through borrowing and lending will not be possible. Risk sharing will therefore more likely take the form of state contingent international income flows. This indicates that we should expect to see the impact of declining home bias on capital income flows most strongly in the lower frequency of the data.[3]

[2] While a number of studies using more recent data also discover a growth in risk sharing and in capital income flows, this increase is relatively modest and remains limited both relative to the growth in international asset gross holdings and to the increase in risk sharing we document here (See Sørensen, Yosha, Wu and Zu (2007) and Kose, Prasad and Terrones (2007) as well as the literature surveyed there).

[3] Another reason for focussing on the longer term in identifying the impact of declining home bias on international capital income flows could be valuation effects (Lane and Milesi Ferretti (2003), Gourinchas and Rey (2007)), i.e. the role that asset prices, and in

A third reason to look for improvements in risk sharing in the lower frequency of the data is that the onset of financial globalization seems to have coincided with a period of major changes in the comovement and volatility of international business cycles. Heathcote and Perri (2004) show that financial integration may actually be associated with lower international business cycle and consumption correlations, suggesting that standard indicators of risk sharing may actually be very susceptible to changes in the underlying structure of business cycle shocks. They provide empirical evidence that the correlation of U.S. consumption growth with consumption growth in the rest of the world has actually decreased since the 1980s. As we have argued in a companion paper (Artis and Hoffmann (2008)), changes in the volatility and synchronization of international business cycles seem to have offset the effect of financial globalization on conventional measures of international risk sharing: the great moderation, the well-documented secular decline in business cycle volatility among industrialised economies, has affected trend output growth by less than it has short-term fluctuations. Since consumption reacts primarily to permanent shocks in income, it appears more volatile relative to output. This impact of the great moderation seems to a large extent to have offset the effect of financial globalization on the volatility of consumption growth conditional on output growth (i.e. on standard formulations of the risk sharing regression in first differences) or on international consumption correlations. In the same mould, Imbs (2006) shows that international trade in goods and assets tends to increase consumption (growth) correlations and output (growth) correlations at the same time, suggesting

particular exchange rate changes, play in the dynamics of foreign asset positions and for the valuation of international income flows. While valuation effects could blur the effect of declining home bias on risk sharing and on capital income flows (either by providing an additional channel of insurance or by acting as a source of shocks), the evidence provided by Gourinchas and Rey (2007) suggests that they matter for external adjustment mainly in the short-run.

that the quantity puzzle (and the question whether there is more risk sharing or actually just less idiosyncratic risk to share) remains unsolved if we focus solely on correlations at the business cycle frequency.

Fourth, a recent influential literature in asset pricing starting with Bansal and Yaron (2004) teaches us that long-term consumption risks are key drivers of expected returns in financial markets. Long-term consumption growth rates seem to be much more highly correlated with asset returns than are short-term fluctuations (see e.g. Parker and Julliard (2005)). These facts further support the notion that we should expect improvements in consumption risk sharing to manifest themselves most strongly at longer horizons.

Recent contributions to which our paper is directly related are Lane and Milesi-Ferretti (2001, 2003) and Sørensen, Wu, Yosha, and Zu (2007).

Lane and Milesi-Ferretti document the virtual explosion in international asset cross-holdings during the 1990s. We find that this phenomenon is a key driving force behind improvements in medium to long term risk sharing through capital income flows.

Sørensen et al. (2007) show that countries with higher shares of foreign assets in their net wealth tend to enjoy better income smoothing through higher international factor income flows. Therefore, the equity home bias and the lack of international consumption risk sharing appear as 'twin puzzles separated at birth'. Our paper maps this idea into a simple theoretical framework. Specifically, we propose a simple model based on Crucini (1999) in which countries can obtain partial income insurance against permanent idiosyncratic shocks through international *ex ante* swaps of claims to each others' output. The parameter that governs the degree of insurance is our theoretical measure of home bias. While countries can further smooth consumption through savings and dissavings, insurance against permanent shocks will only be possible through the capital income flows derived from *ex ante* portfolio diversification.

This integrated framework puts us in a position to make predictions about the quantitative consistency of the decline in home bias – as measured by the growth in international asset cross-holdings –, the growth in international capital income flows and the rise in risk sharing: we find that our theoretical home bias parameter has declined in line with the increase with average OECD asset gross holdings which have grown from 170 to 470 percent of GDP: in particular since the beginning of the 1990s, the fraction of long-term idiosyncratic output risk that gets shared internationally by the average OECD country has increased from less than 10 to more than 30 percent. Our model predicts that virtually all of this increase falls on capital income flows. We find overwhelming confirmation for this prediction in the data.

The remainder of this paper is structured as follows: in the next section we introduce our theoretical framework and use it to motivate the level risk sharing regression. In section three we present our data and estimates of the level risk sharing regressions. In section four we relate the estimated home bias parameter to the patterns of risk sharing, notably of capital income flows. We also show that the longer-term increase in risk sharing through capital income flows that we find in international data is closely related to the growth in international gross asset positions. Section five summarizes and concludes.

2 Income flows, home bias, and consumption risk sharing: an integrated framework

This section presents a general framework that allows us to study the link between consumption risk sharing, portfolio home bias and net capital income flows. Let μ_t^k denote country k's consumption income ratio so that

$$C_t^k = \mu_t^k INC_t^k \tag{1}$$

where C and INC denote per capita values of consumption and income respectively. The only restriction we impose is that $\ln(\mu_t^k)$ should be a stationary variable for the average country in our sample, whereas in our empirical analysis we will explicity allow that the logarithms of consumption and income to be highly persistent, integrated processes. This framework nests virtually all current theories of consumption. For example, the formulation here is consistent with permanent income models, where μ_t^k captures the effect of discounting and uncertainty on consumption, given today's income. We will discuss the interpretation of μ_t^k in more detail below.

Now recall the definition of income (GNI or GNP) from the national accounts: a country's income equals its output plus net claims to output produced in the rest of the world

$$INC_t^k = Y_t^k + NFI_t^k$$

where NFI_t^k is net factor income from abroad, i.e. the country's net claims on flows of foreign output.

In order to link international income flows to the structure of countries' asset portfolios in a tractable manner, we assume that countries trade perpetual claims to their respective output streams. Such assets have first been suggested by Shiller (1993) and we therefore refer to them as Shiller securities. In our setup we follow Crucini (1999) and assume that each country allocates its wealth between a claim to domestic assets and a world mutual fund of foreign Shiller securities. Since income constitutes the dividend from wealth, per capita income must be the weighted average of dividends paid on domestic and foreign assets. The dividends of Shiller securities are just foreign and domestic output, so that per capita income in country k is

$$INC_t^k = \lambda Y_t^* + (1-\lambda) Y_t^k \tag{2}$$

where λ is the (*ex ante* value weighted) share of foreign assets in the country's wealth portfolio. Here, Y_t^k denotes country k per capita output and Y_t^* is the average of per capita outputs across all countries. Note that under these assumptions, net factor income flows are given by

$$NFI_t^k = Y_t^k - INC_t^k = \lambda(Y_t^* - Y_t^k) \tag{3}$$

which captures the idea that countries with higher portfolio shares of foreign assets will achieve more risk sharing through income smoothing.

Perpetual claims to a country's entire output are not currently traded in world financial markets. While certainly stylized, the model is still quite general because existing assets – in particular equity – may at least in part allow countries and regions to replicate the pay-off structure of a portfolio of Shiller securities. In frictionless markets, countries would want to diversify completely, which amounts to selling their national output to the world mutual fund. Hence, under complete diversification, we will have $\lambda = 1$. But claims to a country's entire output would also comprise claims to labour income and other non-tradeable output components. Furthermore, we will expect that frictions in financial and goods markets will drive λ away from unity. The parameter λ is therefore best regarded as a metric of how close observed income flows are to the income flows we would observe if countries or regions could completely diversify any idiosyncratic risk by investing all their wealth in a world portfolio of Shiller securities. One way to think of λ is as the effective degree of diversification of the average country and we refer to it as the 'Shiller portfolio weight' or as 'home bias': the parameter λ tells us what share of a country's income is effectively derived from home and foreign sources.[4]

Portfolio diversification and the ensuing decoupling of income from consumption is the first channel of risk sharing that is nested in our framework. The second is intertemporal smoothing of consumption, e.g. through savings and dissavings. This channel finds its reflection in fluctuations in the consumption-income ratio μ_t^k. Plugging (3) into (1), we obtain

$$C_t^k = \mu_t^k INC_t^k = \mu_t^k \left[\lambda Y_t^* + (1-\lambda) Y_t^k\right] \tag{4}$$

We can think of fluctuations in μ_t^k as capturing an array of country-specific effects such as the rates of return on the country's or region's wealth. For example, in the context of the permanent income hypothesis (PIH), consumption should equal the permanent component of income, INC_t^{kP}, so that $\mu_t^k = INC_t^{kP}/INC_t^k$ reflects time variation in the discount factor and / or the presence of temporary components in income (see e.g. Campbell and Mankiw (1989)). While such fluctuations in μ_t^k may be an important source of risk sharing (consumption smoothing) at short to medium horizons, we

[4]The parameter λ can also be thought of as a long-run partial insurance parameter in the sense of Heathcote, Storesletten and Violante (2007).

assume that μ_t^k is stationary, so that consumption and income move together in the long-run. Hence, in the long-run any risk sharing in this model can only come from income smoothing, suggesting that portfolio diversificiation and long-term risk sharing through capital income flows go hand in hand. Indeed, as we show next, $1 - \lambda$ also measures the fraction of variability in long-term relative output movements that systematically spills over into consumption – it is a measure of long-run risk sharing.

2.1 A risk sharing regression in levels

To obtain an estimation equation for λ, we first re-write equation (4) by dividing by world income. Let μ_t^* denote the world consumption-income ratio and use the fact that world per capita income equals world per capita output, to obtain:

$$\frac{C_t^k}{C_t^*} = \frac{\mu_t^k}{\mu_t^*}\left[\lambda + (1-\lambda)\frac{Y_t^k}{Y_t^*}\right] \tag{5}$$

This relation can straightforwardly be log-linearized around $Y_t^k/Y_t^* = 1$ to obtain

$$c_t^k - c_t^* = (1-\lambda)\left[y_t^k - y_t^*\right] + \phi_t^k \tag{6}$$

where lower-case letters denote natural logarithms and $\phi_t^k = \log\mu_t^k - \log\mu_t^*$ denotes the relative log consumption-income ratio.

This equation is reminiscent of the equations estimated in Mace (1991), Asdrubali, Sørensen and Yosha (1996), Cochrane (1991) or Crucini (1999). It relates relative consumption to relative output. Under full risk sharing, the coefficient on relative (permanent) output should be zero. The decisive difference vis-à-vis the earlier literature is that equation (6) relates relative log-levels whereas earlier implementations were formulated in differences.

Equation (6) lets us estimate λ based on consumption and output data. The estimate of λ, can then be used to predict international income (as opposed to output) data simply from equation (2) which can then be compared with actual international income data. Specifically, to the extent that financial globalization is associated with improved long-term risk sharing, i.e. a decline in $1 - \lambda$, our model suggests that this improvement should be explained by an increase in international net factor income flows, as predicted by the relation $NFI_t^k = \lambda(Y_t^* - Y_t^k)$. Furthermore, the dual interpretation of $1 - \lambda$ as a measure of home bias, also suggests that a decline in λ should be associated with a growth in international investment positions. We explore both implications below.

In our empirical analysis we estimate (6) as a panel relation. Our setup allows for a high degree of unobserved heterogeneity across regions and countries through the time-varying, country-specific term ϕ_t^k. While ϕ_t^k can capture a wide array of different influences, the requirement that a country fulfill its intertemporal budget constraint imposes an important restriction: the process ϕ_t^k reflects fluctuations in relative consumption-income ratios and intertemporal budget balance requires that, in the long run, a country's consumption will have to correspond to its income. While, even in the long-run, consumption and income may not literally be identical, e.g. because the country can systematically reap capital gains from foreign assets or because precautionary motives induce the country's residents to consume less than their actual income, we should still expect the long-run mean of μ_t^k and therefore also of ϕ_t^k to be constant, so that ϕ_t^k should be stationary. This stationarity of ϕ_t^k implies that (6) constitutes a panel cointegrating relationship. We discuss the econometric issues involved in the estimation of this relationship below.

We highlight the empirical advantages of our approach vis-à-vis the differenced versions of the risk sharing regression (6). Differenced specifications of the risk sharing regression or correlation-basesd measures of risk sharing (see e.g. the results in Artis and Hoffmann (2008) and Imbs (2006), Bai and Zhang (2005)) often do not appear to pick up an increase in risk sharing over time. Heathcote and Perri (2004) even report that international consumption correlations have decreased for the U.S. Conversely, our results here clearly show an increase in risk sharing (a drop in the estimated coefficient) when we use the levels regression (6).

Why does the level regression pick up an increase in risk sharing that the differenced regression does not seem to register? Recall that the term ϕ_t^k is the relative consumption-income ratio. According to standard consumption theory, fluctuations in the consumption-income ratio should be a good indicator of future expected discount rates or of future income changes. For example, in a permanent income model, if income changes are serially correlated, changes in ϕ_t^k will be correlated with changes in income and therefore also with $\Delta y - \Delta y^*$. Specifically, in Artis and Hoffmann (2008) we show that if consumption follows permanent income, the coefficient b in the differenced risk sharing regression

$$\Delta c_t^k - \Delta c_t^* = b\,[\Delta y_t - \Delta y_t^*] + \varepsilon_t^k$$

can be written as

$$b = \delta(1 - \lambda)$$

where $\delta = var(\Delta y_t^P - \Delta y_t^{*P})/var(\Delta y_t - \Delta y_t^*)$ is what we call the long-run variance ratio, i.e. the ratio between the variability in the permanent component of relative output growth, $\Delta y_t^P - \Delta y_t^{*P}$, and the variability of output growth overall. Unless $\delta = 1$, the coefficient in the differenced regression will be a biased estimate of $(1 - \lambda)$ and the bias will be a function of the particular nature of shocks, as captured by δ. In particular, if the structure of cyclical shocks changes, affecting δ, the effect of financial integration on consumption correlations or the differenced risk sharing regression may get blurred.[5] As is well-known, in a cointegrating regression such as (6), this simultaneity is not a problem because the joint non-stationary dynamics in relative output and consumption tends to dominate the stationary term ϕ_t^k in the estimation. Thus, we should expect the level risk sharing regression to pick up longer-term trends in risk sharing more robustly.

3 Empirical Implementation

3.1 Data

We examine annual data for 23 OECD countries ranging from 1960 to 2004. Consumption, GDP and GNP (income) and population are from the Penn World Table, release 6.2 (PWT 6.12.) by Heston, Summers and Aten (2006). All data are in constant (2000) prices. The countries included in our estimation are:

1. Canada, 2. the United States, 3. Japan, 4. Austria, 5. Belgium, 6. Denmark , 7. Finland, 8. France, 9. Germany (West), 10. Greece, 11. Iceland, 12. Ireland, 13. Italy, 14. Luxemburg, 15. Netherlands, 16.

[5] The quantitative-theoretical considerations in Heathcote and Perri (2004) strongly suggest that the nature of international shocks has changed since the beginning of the 1980s. Our companion paper (Artis and Hoffmann (2008)) explores the impact of this 'great moderation' on consumption correlations and the differenced regressions commonly used in the risk sharing literature. In fact, financial integration itself could be responsible for such changes (see Imbs (2006)).

Norway, 17. Portugal, 18. Spain, 19. Sweden, 20. Switzerland, 21. United Kingdom, 22. Australia, 23. New Zealand.

We also obtain results for U.S. state level data. While interesting in their own right — level risk sharing regressions as we propose them here have not previously been estimated on regional data — the regional results from a financially highly integrated economy such as the U.S. will provide a natural benchmark against which we can evaluate our international results. The US-data set is the one used in the seminal paper by Asdrubali, Sørensen and Yosha (1996) where it is also described in detail. Output and income are measured by gross-state product and state-level personal (disposable) income data respectively, both from the Bureau of Economic Analysis (BEA) regional economic accounts. Consumption data at the state level are not available. We follow Asdrubali, Sørensen and Yosha (1996) and virtually the entire literature on regional risk sharing in the U.S. by using retail sales data. State-level retail sales data are re-scaled by the share of retail sales in aggregate (US-wide) consumption to obtain measures of state level consumption. Following the general practice in the US regional business cycle literature and drop Washington D.C. from the sample. All data are deflated by the US-wide consumption price index and range from 1960 to 1990.

We express all data in per capita terms. Rest of the World (RoW) aggregates are the US- or OECD-wide average per capita values. Population data are from the BEA and PWT respectively.

Over the sample period covered by our international data set, international financial markets have become increasingly liberalized. To take account of this change, we will report results obtained for two subperiods: the first covers the period 1960-1990, the second covers 1990-2004. The results we obtain from the first sub-period can be compared directly to others in the literature (the studies by Sørensen and Yosha (1998) and Crucini (1999) cover the same period), while the results from the second sub-period should provide insights into the effects of the dramatic growth in gross international asset positions since the beginning of the 1990s.

3.2 Estimating the home bias parameter

We now turn to estimating the parameter $1 - \lambda$ based on the level risk sharing regression (6). In the light of our previous discussion, we treat equation (6) as a cointegrating relationship.[6]

This relationship can, in principle, be estimated consistently by OLS. However, OLS may suffer from second-order bias due to potential simultaneity and serial correlation of the errors. Phillips and Moon (1999) therefore advocate a panel version of the fully modified least squares (FMLS) method. Since the FMLS estimator is semiparametric, it may, however, be imperfectly suited to relatively small samples. The panel dynamic OLS (PDOLS) estimator suggested by Mark and Sul (2003) may be preferable in this case. We therefore conduct all our analyses here based on the panel OLS and the panel dynamic OLS estimator.

The panel dynamic OLS estimator accounts for serial correlation and potential simultaneity by including leads and lags of the differences of the right hand side variables. We experimented with various leads and lags and found our results to be very robust across specifications. All results were also very similar to those obtained from plain panel OLS estimates. The

[6] Whether or not (6) constitutes a cointegrating relationship is without consequence for the point estimates of $(1 - \lambda)$. As Phillips and Moon (1999) note, the coefficient of a non-stationary panel regression is meaningful even if there is no cointegration between the variables. The cointegration tests we report below do, however, support the view that (6) is a cointegrating relation.

PDOLS parameter estimates in table 1 are based on one lead and lag for each country which we found sufficient to capture serial dependence in our annual data. To ensure that we can really treat (6) as a panel cointegrating relation, we perform Pedroni's (2004) group mean test for panel cointegration on the residuals of the country-wise PDOLS-regressions. These tests strongly support rejections of the null of no cointegration in all samples and subperiods. They are reported as memorandum items at the bottom of table 1.

In our estimation, we assume that the relative consumption-income ratio ϕ_t^k can be decomposed into a common-time-specific component τ_t, a country- or region-specific fixed effect ϕ_k and into a country-specific mean zero process u_t^k so that

$$\phi_t^k = \tau_t + \phi_k + u_t^k$$

The modelling of the term ϕ_t^k and in particular the treatment of the region- or country-specific fixed effect ϕ_k affects the interpretation of the estimated coefficients as measures of risk sharing and international diversification: the pooled estimate, similar to a between-estimator, emphasizes the cross–sectional dimension of the data by abstracting from all the within-country or region variation in ϕ_t^k. For the between estimator to be consistent, the cross-sectional covariance between ϕ^k and relative output has to be zero. Since ϕ^k is the mean relative consumption income ratio, fluctuations in which can help shield consumption against fluctuations in income (and, given λ: also in output), this assumption amounts to saying that intertemporal consumption smoothing as the second channel of risk sharing that is nested into our framework should be absent. In our model, this needs to be the case only in the very long run. We therefore associate the pooled estimate with our long-run risk sharing and portfolio home bias measure $1 - \lambda$.

Clearly, cross-sectional variation in ϕ_k could, at least to some extent, capture unobserved cross-country differences that possibly originated in the very remote past and that are unrelated to risk sharing. We therefore also present fixed effect regressions. The fixed effect estimate, that we denote β_U, accentuates the impact of of departures of relative consumption-income ratios from their sample mean.[7] The cross-sectional covariance of $\phi_t^k - \phi^k$ with the departure of relative output from its long-run (sample) mean is a measure of the contribution of intertemporal consumption smoothing to risk sharing. As argued above, we do not expect fluctuations in the consumption-income ratio to contribute to risk sharing in the very long-run, though. We therefore interpret β_U as a measure of risk sharing at medium horizons, *given* the long-run portfolio home bias captured by $1 - \lambda$.

We present the results from pooled and fixed effect level risk sharing regressions in table 1.[8] Our findings carry a clear message: for US federal states, we find a home bias, $1 - \lambda$, of around 50 percent. Virtually the same coefficient is obtained once we control for fixed effects, so that $\beta_u \approx 1 - \lambda$. In international data, based on the pooled estimate, we detect a home bias of over 90 percent in the 1960-90 period. For the later (i.e. the globalization) period, estimates of $(1 - \lambda)$ are around 0.75. This is considerably lower than in the 1960-90 period and the difference does appear to be significant. Turning to the fixed effect estimates, the increase in international risk sharing in the post-1990 period appears even more marked: for the 1960-90 period we

[7] The fixed effect panel OLS estimator can be written as the time average of the cross-sectional regressions of the previously demeaned variables (see e.g. Asdrubali, Sørensen and Yosha (1996)).

[8] In both cases, like in all panel regressions reported in the remainder of the paper, we control for time-fixed effects.

now find virtually no risk sharing ($\beta_U = 0.98$), whereas for the globalization period the corresponding value is around $\beta_U = 0.65$. Note that the choice of estimation method (OLS vs. PDOLS) has practically no effect on the results.[9]

Our results suggest that there is a lack of risk sharing in international data in both subperiods, but even at the regional level we find that U.S. citizens own a disproportionate share of the claims to output of the federal state in which they live – around 50 percent of regional wealth is held in assets in the own region. This result provides a perspective on the relative size of intra- and international home bias: by measuring the effective degree of financial integration, we also take account of those components of a nation's or region's output risk that are not traded in financial markets: the equity of small firms or companies is most likely not traded across countries or regions nor are claims to the labour share of national or regional outputs. Our estimates seem to reflect this.[10] Against this background, our estimate in table 1 of the increase in international portfolio diversification during the 1990s – a drop in the value of $1 - \lambda$ from 0.91 to 0.74 for the pooled estimate – amounts to a dramatic increase in international risk sharing: if our empirical measure of home bias among U.S. states (~ 0.5) is taken as a benchmark,

[9]This result is informative about the empirical relevance of theoretical scenarios in which consumption and output are incidentally correlated for reasons that are unrelated to financial markt incompleteness, for example, due to preferences that are non-separable in consumption and leisure as in Backus, Kehoe and Kydland (1992). The fact that we do not find major differences between the PDOLS - which implicitly controls for this incidental correlation – and the panel OLS estimates supports the notion that such non-separabilities are not likely to have an important effect on risk sharing regressions. This is in line with Backus, Kehoe and Kydland's own conclusion that the non-separability between consumption and leisure cannot quantitatively resolve the consumption correlation puzzle.

[10]Note that risk sharing among U.S. federal states will also partly be achieved through net fiscal transfers. Asdrubali et al. (1996) show that this channel accounts for roundabout 10-15% of the risk sharing achieved among U.S. federal states at business cycle frequencies. Becker and Hoffmann (2006) report similar results for the long-run. Given these findings, the increase in risk sharing through private capital markets at the international level – where fiscal transfers are virtually absent – would appear relatively even more important.

then around 40 percent $((0.91 - 0.75) / (0.91 - 0.47) = 0.17/0.43 \approx 0.4)$ of the international home bias (relative to regional home bias) has vanished in the years after 1990. Based on the fixed effect estimate, the increase in risk sharing would appear even more dramatic; the increase in risk sharing in international data is not only statistically significant but also economically important.

In Table 2 we provide a number of checks that illustrate that the increase in international risk sharing is indeed a robust feature of the data. First, we run our regressions on the 1975-1990 subperiod which is of equal length as the globalization period 1990-2004. Secondly, to rule out that our findings are unduly affected by small but financially very open economies such as Switzerland, Ireland, Luxemburg and Iceland, we experiment with excluding these countries from the sample. For comparison, we also run the regressions for a sample of 11 small open economies only. Third, we further address the issue that relative consumption and output levels may be nonstationary variables. As we have argued before, spurious regression problems are are not much of an issue in non-stationary panels (see Phillips and Moon (1999)). In addition, our results in table 1 suggest that the risk sharing regression identifies a panel cointegrating relationship. Nonetheless we provide additional evidence for the robustness of our general conlusions by differencing the data at horizons of three and five years.[11] As is apparent from the table, the gist of our results remains unaffected by all of these exercises: the more we emphasize the lower frequency of the data, the more clearly we see a decline in the estimated coefficients over time – there is a clear increase in international risk sharing. The effect for the pooled

[11] Of course, using data that has been differenced at longer horizons to emphasize the low-frequency interaction willl lead to a loss of a substantial number of oberservations in the annual samples we are considering here, making the method possibly less efficient than our level regressions.

estimate – that, as we argued, is associated with the long-run home bias parameter $(1-\lambda)$ – seems particularly clearcut.

4 Patterns of risk sharing and international asset positions

Since the early 1990s average gross foreign asset positions of OECD countries have grown dramatically (Lane and Milesi-Ferretti (2001, 2005, 2006)). In this section, we demonstrate that the main channel through which the increase in international consumption risk sharing has occured is through an increase in international capital income flows and that both developments – better risk sharing and bigger role for international capital income flows – can directly be traced to this growth in gross international investment positions. – very much as our theoretical framework would suggest.

4.1 Channels of risk sharing

Our theoretical framework nests two channels of risk sharing. The first are capital income flows, that we associated with the (logarithmic) ratio between a country's or region's output and income. The covariation of this ratio with fluctuations in relative output indicate to which extent a country can smooth its income in the fac of output shocks. Building on Asdrubali, Sorensen and Yosha (1996), we measure the importance of this channel through regressions of the form

$$\left[y_t^k - y_t^*\right] - \left[inc_t^k - inc_t^*\right] = \theta_k^K + \beta_K \left[y_t^k - y_t^*\right] + v_t^K \qquad (7)$$

where 'inc' denotes the logarithm of income and θ^K is a country or region fixed effect. We think of international or interregional capital income flows as being derived from *ex ante* diversification of countries' or regions' portfolios. We call the channel the capital income flow, income smoothing or *ex ante* channel.

Secondly, countries or regions can further smooth consumption relative to income through intertemporal asset trade. Variation in the consumption-income ratio[12] captures to what extent a country or region manages to decouple its consumption. In our model, this is represented through time-variation in ϕ_t^k. We refer to this second channel as the *ex post* or intertemporal channel or as consumption smoothing. We measure its contribution to risk sharing through regressions of the form

$$\left[inc_t^k - inc_t^*\right] - \left[c_t^k - c_t^*\right] = \theta_k^C + \beta_C \left[y_t^k - y_t^*\right] + v_t^C \qquad (8)$$

Note that $\left[inc_t^k - inc_t^*\right] - \left[c_t^k - c_t^*\right] = -\phi_t^k$.

Both the *ex ante* and *ex post* channels together account for the entirety of consumption risk sharing, so that

$$\beta_K + \beta_C = 1 - \beta_U$$

and β_U is the unsmoothed share of idiosyncratic risk, which can be obtained from the regression

$$\left[c_t^k - c_t^*\right] = \theta_k^U + \beta_U \left[y_t^k - y_t^*\right] + v_t^U$$

which is exactly the fixed effect level risk sharing regression that we have estimated in the previous section so that $\theta_k^U = \phi^k$.

We call a particular combination of β_K and β_C a *pattern* of risk sharing. It is apparent that, while the total extent of risk sharing $1 - \beta_U$ does not

[12]Note that the (relative) consumption-income ratio corresponds to ϕ_t^k in (1) above.

depend on income, the dynamics of income flows is key for the pattern of risk sharing: if there is no *ex ante* diversification of a country's or region's asset portfolio, we would expect income and output to be the same and $\beta_K = 0$. Hence, any risk sharing would only occur *ex post* as consumption smoothing. Conversely, the same amount of overall risk sharing $1-\beta_U$ could be achieved *ex ante* through income flows alone, with consumption equalling income.

Thus, given β_U, the dynamics of income determine the pattern of risk sharing. Since, our model predicts that income is

$$INC_t^k = \lambda Y_t^* + (1-\lambda)Y_t^k \tag{9}$$

the values of β_K and β_C depend directly on λ. Specifically, in the very long run, we would expect that $\beta_C = 0$ and that therefore $1 - \beta_K = \beta_u = 1 - \lambda$. Clearly, in a given finite sample, this will not necessarily be the case and our long-run estimate of risk sharing and home bias, λ, and of its medium-run counterpart β_u will generally differ and we will therefore generally have $\beta_K \neq \lambda$.[13]

Using our estimate of λ, we generate artificial income data according to (9), based on actual values of GDP for Y^* and Y^k. We then compare predicted income to actual GNP (national income) and state-level income. To conduct this comparison, we estimate the panel equations (7) and (8) using real-world data for y and c, but once based on actual income data and a second time using our artificial income data. We emphasize that this exercise imposes a strong restriction on our model because λ has been estimated from consumption and *output* (GDP or gross state product) data alone,

[13]Specifically, as we have done in the previous section, we associate β_u (and also β_K and β_C)with the fixed effect estimate and λ with the pooled estimate.

without ever making use of *income* (GNP or state level personal income) data. Given output and consumption, the correlation of income with output and its variability determine the pattern of international risk sharing, β_K and β_C, so that running the regressions (7) and (8) on artificial data checks whether our estimate of λ allows us to generate patterns of risk sharing that we see in the data. In view of the fact that variance decompositions in the mould of Asdrubali, Sørensen and Yosha (1996) have become a standard metric in the empirical risk sharing literature, this exercise appears as a natural test of our model.

For the U.S., the real world income data we use for this exercise is state-level personal disposable income.[14] In international data, we measure income through gross national product (GNP). The Penn World Tables contain GNP only after 1970. We therefore limit our analysis of international data from now on to the post-1970 period. This does not affect the interpretation of our results since our estimates of $1-\lambda$ for the 1970-90 period are virtually identical to those obtained for the 1960-90 period.

For the generation of the artificial income data, we use a value for $1-\lambda$ of 0.48 for the United States. For the OECD countries, we set $1-\lambda$ to 0.91 in the 1970-90 period and to 0.74 in the period 1990-2004. These values correspond to our estimates of $(1-\lambda)$ in table 1, based on the PDOLS procedure in the model without fixed effects. In the bottom panel (III) of Table 2, we report the average across countries and regions of the correlations between artificial and actual idiosyncratic income growth rates. For the average U.S. state the correlation between actual and fitted income growth is 0.77. At the international level, the model performs even better: here, it generates an average correlation of 0.94 in the 1970-90 period and of 0.85

[14] For the U.S., we compare the artificial data generated from (9) with disposable income because λ may in part reflect permanent fiscal transfers.

in 1990-2004.

We estimate equations (7) and (8) using panel dynamic OLS with one lead and lag, controlling for fixed effects and for common time-specific variation. The upper panel (I) in Table 3 reports the results based on both predicted and actual income data.

For the U.S., the predicted pattern of risk sharing is almost identical to that found in the data. In addition – very much as our model would predict for the very long-run – the values of β_K and λ are virtually the same. The estimates clearly suggest that most risk sharing in the U.S. takes place through capital income flows. While there is some consumption smoothing *ex post*, this appears overall insignificant. This confirms our previous expectation that – at least in the longer term emphasized in our theoretical and empirical framework – better risk sharing will have to be associated with larger capital income flows.

Though our model would somewhat overpredict the amount of *ex ante* risk sharing in the 70-90 period (the predicted β_K is 0.08 vs. 0.01 in the data), it remains true that the overall pattern of risk sharing is very much in line with what our model would predict: these are the results for the period *before* the huge internationalisation of asset ownership and based on our model we would expect that the huge portfolio home bias that prevailed at that time would be reflected in low capital income flows.

The picture changes completely once we turn to the globalization period: now there is a considerable amount of risk sharing in international data, roundabout one quarter to 30 percent of idiosyncratic output risk gets shared *ex ante*. There is also a moderate increase in *ex post* consumption smoothing. But this effect is much more subdued and the *ex post* channel appears only marginally significant. Again, the regressions based on predicted income get very close to the real-world pattern of risk sharing.

These results provide a complementary perspective to an important literature that has shown that international risk sharing is lower than in U.S. data mainly because international capital income flows do not contribute to risk sharing (Sørensen and Yosha (1998), Becker and Hoffmann (2006)). Lane (2001) concludes that international investment income flows have practically no bearing on risk sharing internationally. These analyses have typically focused on business cycle frequencies of the data by looking at versions of the regressions (7) and (8) that are formulated in first differences. In panel II of Table 3 we estimate such differenced regressions, again for actual and predicted income flows.

The differenced regressions based on actual income clearly corroborate the findings in the studies referenced above: The lack of international consumption risk sharing frequently is mainly a lack of *ex ante* income smoothing: for the 1960-90 period we find roundabout 40 percent income smoothing for the U.S. and virtually none in international data. We find a small rise in international consumption risk sharing in the 1990-2004 period, with the sum of the coefficients of the differenced channel regressions, $\beta_K^\Delta + \beta_C^\Delta$, increasing from 0.27 to 0.33. But this difference seems tiny and appears insignificant. Furthermore, even in the globalization period, only a small share of an increase would be explained by capital income flows; β_K^Δ increases from 0.03 to 0.06 and, again, the increase is insignificant.

We are primarily concerned with the long-term relation between income flows, home bias and consumption risk sharing. Still, it is interesting to see that our model is able to match the business cycle frequency patterns of risk sharing among U.S. federal states very well. In international data, given the huge decline in $1 - \lambda$, we should expect our model to predict a lot more income smoothing also in the short run. This is exactly what we see, but it is at odds with the estimates for β_K^Δ and β_C^Δ from actual data that, as we have just seen, have barely budged. These results show that at short horizons the increase in international risk sharing *a)* appears rather limited and is, *b)* difficult to associate with a marked shift in risk sharing patterns towards higher international income flows. At longer horizons, however, this increase is much more readily apparent and the patterns of international capital income flows line up with the predictions from our simple model.

4.2 The increase in international risk sharing and the growth in gross international asset positions

In this subsection, we show that the decline in our home bias measure $1-\lambda$ is closely linked to the internationalisation of asset ownership. As our measure of international portfolio diversification we use gross asset positions, the sum of assets and liabilities, relative to GDP. This choice is based on some *a priori* theoretical and empirical considerations. First, Obstfeld (2004) distinguishes between two motives for asset trade: intertemporal or development asset trade, which is reflected in net investment positions and diversification asset trade, which he associates with gross asset positions. Our interest in this paper is clearly in the risk sharing or diversification aspect of asset trade. Secondly, measuring home bias through gross asset positions is consistent with our model in which the average country swaps a share λ of claims to its own output for a share λ of claims to world average output. Hence, λ is a measure of the cross-sectional average of country's gross asset positions. Finally, the focus on gross asset positions can also be justified at an empirical level: Lane and Milesi-Ferretti (2003) note that there has been a virtual explosion in cross-holdings of assets whereas net positions have remained quite stable.

To obtain a detailed time profile of the decrease in home bias we now let λ vary from period to period. We do so in two ways: first, by running a sequence of cross-sectional versions of our level risk sharing regression:

$$c_{kt} - c_t^* = (1 - \lambda_t) \left[y_t^k - y_t^* \right] + \tau_t + u_{kt} \tag{10}$$

Secondly, in order to capture the impact of increased gross-asset holdings over time, we modify our panel level risk sharing regression to take account of the internationalisation of average gross asset positions. Specifically, we parametrize λ as

$$\left(1 - \lambda_t^{GFA}\right) = \kappa_0 + \kappa_1 GFA_t \tag{11}$$

where GFA_t is the average (across countries) gross foreign asset position relative to GDP:

$$\overline{GFA}_t = \frac{1}{K} \sum_{k=1}^{K} GFA_t^k \text{ and } GFA_t^k = \frac{A_t^k + L_t^k}{Y_t^k}$$

Here, K denotes the number of countries, A_t^k is assets and L_t^k liabilities and the bar denotes the cross-sectional mean. Our data source is the March 2006 release of the external wealth of nations data set by Philip Lane and Gianmaria Milesi-Ferretti (2006). To obtain estimates of κ_0 and κ_1, we plug (11) into our panel level risk sharing regression and estimate the ensuing regression with interaction terms by panel dynamic OLS. Based on these estimates, we can then obtain a second time series of $1 - \lambda_t$ from (11). We refer to this alternative estimate of λ as λ^{GFA}.

We obtain these two time-varying measures of λ for the period from 1975 to 2004. Using panel dynamic OLS, we estimate

$$\left(1-\lambda_t^{GFA}\right) = 1.01 - 0.1\overline{GFA}_t$$

with t-statistics on κ_0 and κ_1 of 19.97 and 2.94 respectively. There is a statistically significant link betweeen average gross foreign asset positions and risk sharing. Our estimate of κ_0 is virtually unity, suggesting that cross-holdings of assets seem to account for all the consumption risk sharing we see in the data. Finally, it is interesting to appreciate the magnitude of the coefficient κ_1: increasing average gross foreign asset holdings by 100 percent of GDP will increase consumption risk sharing by roundabout 10 percentage points.

In Figure (1) we plot both the estimated coefficients $(1-\lambda_t)$ from the sequence of cross-sectional regressions in (10) as well as the measure based on the parametrization (11), $1-\lambda_t^{GFA}$. Though obtained from different approaches, the two measures both clearly indicate an increase in risk sharing that is of a very similar magnitude. We obtain standard errors for $1-\lambda_t$ from a jackknife procedure (Efron (1982)), in which we re-estimate the sequence $1-\lambda_t$ (for $t = 1975...2004$) 23 times, dropping one country from the sample at a time. The tightness of the two standard error bands thus obtained also demonstrates that our conclusions concerning the extent of long-term risk sharing and its increase during the recent globalisation period are robust with respect to the inclusion or exclusion of individual countries in our sample. In addition, over most of the sample period, the GFA-based measure is generally well within the bands, suggesting that it is almost statistically indistinguishable from $1-\lambda_t$.[15] We find this particularly remarkable since

[15] Only towards the end of the sample period, the two measures diverge somewhat, as the decline in $1-\lambda_t$ seems to flatten out for a couple of years. A potential explanation is that the global stock market decline after 2000 has lowered the value of equity and other financial assets relative to real assets, notably human wealth and housing. Since holdings of these assets are not internationally diversified, international risk sharing may well have

we have not been using any information about asset holdings in estimating the sequence of cross-sectional regressions (10). These findings suggest that the decline in our home bias measure and the associated rise in risk sharing are indeed intimately linked. However, the two variables GFA and λ are trending and in order to avoid spurious conclusions, we also explore the possibility that some other trend, unrelated to the growth in GFA, might be driving our results. To this end, we run regressions in which we also include a linear trend term into the specification of λ_t^{GFA} so that

$$1 - \lambda_t^{GFA} = \kappa_0 + \kappa_1 \overline{GFA}_t + \kappa_2 t \tag{12}$$

Table 4, panel I juxtaposes the results to those obtained from the previous specification for λ^{GFA}. Over the whole sample period (1975-2004), including the linear trend term makes both the trend and GFA insignificant, a pattern that could reflect the collinearity of the two variables. We also find that, both GFA and the linear trend are insignficant before 1990. However, GFA is clearly significant after 1990, even once we control for a linear trend. This confirms our previous conjecture: as noted by Lane and Milesi-Ferretti, international asset cross-holdings really started to take off only during the last decade of the previous century and we should therefore expect to see the impact of financial integration on risk sharing mainly after 1990.

To make the point that it is the growth in international cross-holdings of assets that drives the increase in international risk sharing even more tightly, we perform the following exercise: At each point in time, we sort countries

decreased temporarily only to pick up again in 2003 with the recovery of international stock markets. Since $1 - \lambda_t^{GFA}$ is a function of gross foreign asset positions only – which have continued to grow relative to GDP – it does not detect this temporary decrease in risk sharing.

by the size of their gross foreign asset position relative to GDP, i.e. by GFA. We then form synthetic panel groups of 'portfolios' of countries based on the quartile in which a country is at a given time in the distribution of gross foreign asset positions.[16] We then run our level risk sharing regressions above based on the panel of observations from the lowest and highest quartile. The results are reported in panel II of Table 4. As is clearly apparent, the countries with the highest GFA position at a given point in time clearly enjoy a higher of level of risk sharing. This is true for the entire period as well as for the two subperiods, even though it is clearly more pronounced again after 1990.

Finally, we turn to exploring how the internationalisation of asset positions has affected the patterns of risk sharing. A salient feature of the results in table 2 that distinguishes our findings from most of the earlier literature is that the increase in international risk sharing is strongly associated with a more important role for international income flows. We show that this changing pattern is also directly related to the growth in GFA, in analogy to (12) we parametrize the pattern of medium term risk sharing, β_K and β_C, as a function of international asset positions. To exploit information concerning the cross-sectional variation in asset holdings, we also let β_K and β_C vary across countries by using the time t - country k realization of

[16] E.g. in our sample of 23 OECD countries, the synthetic panel built on the lowest quartile contains the artificial time series of consumption and output for the six countries with the lowest GFA-positions $\widetilde{c}_t^l$ and $\widetilde{y}_t^l$ for $l = 1..6$ where $\left\{\widetilde{c}_t^l, \widetilde{y}_t^l\right\} = \left\{c_t^k - c_t^*,\ y_t^k - y_t^* \,|k \text{ such that country } k \text{ has the } l\text{-th lowest GFA-position at time } t\right\}$.

Note that this procedure – quite similar to the practice in the asset pricing literature of sorting assets into portfolios with time-varying composition – should be more powerful in identifying the effects of international investment positions on risk sharing than a conventional sample split (based e.g. on the sample period average gross foreign asset position): because countries change their relative positions and therefore their quantile groups, it is unlikely that slow moving fundamental but unobserved country differences that could be correlated with large *average* cross holdings rather than gross holdings themselves are responsible for the higher level of risk sharing that we identify for countries with high GFA positions. We note, however, that results from a simple sample split into high and low GFA countries give very similar results.

GFA instead of its cross-sectional, mean, so that

$$\beta_{Xt}^k = \beta_{0X} + \beta_{1X} GFA_t^k + \beta_{2X} t$$

and where X stands for K and C in turn. Plugging this specification into our channel regressions, we again obtain a set of regressions with interaction terms between relative output and GFA and t respectively. Again, we also consider a specification without a linear trend.

The impact of certain asset categories on the patterns of risk sharing may not necessarily be uniform. In particular, we would expect that equity assets and liabilities to have a more direct impact on state-contingent international capital income flows than debt instruments. We therefore also express the pattern of risk sharing as a function of the growth in international gross equity positions and of debt holdings, parametrizing

$$\begin{aligned}
\beta_{Xt}^k &= \beta_{0X} + \beta_{1X} GFE_t^k + \beta_{2X} t \\
\beta_{Xt}^k &= \beta_{0X} + \beta_{1X} GFD_t^k + \beta_{2X} t
\end{aligned}$$

where GFE_t^k (GFD_t^k) is the sum of equity (debt) assets and liabilities divided by GDP

Table 5, panel I reports the results for these interaction term regressions. In panel II we again report results from synthetic panel regressions, where the panel groups are formed on the basis of sorting countries at any point in time according to their GFA-, GFE and GFD-positions respectively. The analysis in table 5 is limited to the 1990-2004 period, for which we have uninterrupted data by asset category for all 23 countries. The first two rows are for total assets, rows 3 and 4 for equity and rows 5 and 6 for debt assets.

The picture emerging from table 5 is clearly that the growth in international cross-holdings tends to increase risk sharing – the sum $\beta_{K1} + \beta_{C1}$ is positive and significant for all three categories of assets. Most interestingly however, the growth in GFA, GFE and GFD has a particularly strong effect on β_K, i.e. the contribution of capital income flows to risk sharing. In line with prior expectations, this effect is particularly strong for cross-holdings of equity assets: equity is more likely to deliver state contingent dividend payments whereas debt assets will not generally provide such income insurance. Conversely, debt assets have a relatively stronger impact on risk sharing through ex post accumulation or decumulation of foreign assets (β_C).

5 Summary and Conclusion

A key finding of our paper is that consumption risk sharing among OECD countries has improved dramatically over the last decade. This finding is what one should expect in a world where the barriers to international capital flows have virtually been removed and it ties in with the bulk of empirical evidence that suggests that international cross-holdings of claims to capital have grown considerably. Still, the literature so far has found it relatively difficult to document that higher capital mobility actually *does* find its reflection in better international consumption risk sharing.

As we have argued, there are a number of theoretical and empirical reasons to believe that improvements in risk sharing show up first and most robustly in the lower frequency of the data. In our analysis, we have emphasized these lower frequencies by making use of the information implicit in the relative levels of output and consumption. Unlike many earlier econometric specifications that have focused on the business cycle link between consumption and output, this has allowed us to document that risk sharing has indeed increased considerably among industrialised economies. By the end of our sample period, more than 30 percent of the long-term idiosyncratic risk faced by the average OECD country gets shared internationally. The main channel through which these improvements in international risk sharing seem to have come about is through a larger contribution of international capital income flows to the smoothing of national income rather than through direct smoothing of consumption through savings and dissavings.

We discussed our findings in the framework of simple theoretical model that captures the idea that countries can obtain partial income insurance against permanent idiosyncratic shocks through ex ante swaps of claims to each others' output. The parameter that governs the degree of long-term partial insurance is our theoretical measure of home bias. Out of observed income, countries can then – over the short and medium term – further smooth consumption through savings and dissavings. However, in this model, all long-run insurance must ultimately come from capital income flows. This is what we observe in the data. In fact, we find that the increase in international capital income flows that we observe in the data since the 1990s is not only qualitatively but also quantitatively consistent with this simple model. In particular, it is possible to estimate our parameter of long-term partial insurance – our theoretical measure of 'home bias' – from nothing but data on output and consumption. Once this parameter is known, we can then generate international capital income flow data. As we show, our model replicates the patterns of international risk sharing and capital income flows observed in the data with remarkable precision.

Finally, we demonstrated that both the rise in risk sharing and the increase in international capital income flow can indeed be directly linked to the growth in international asset gross-holdings. We find that higher gross asset positions lead to considerably more long-run insurance and that capital income flows play a more pronounced role as a channel of risk sharing for those countries that have large cross-holdings of asset, notably of equities.

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Table 1: The increase in long-run risk sharing

Panel I: Pooled Regressions (w/o fixed effects)

$$(c_t^k - c_t^*) = const + \tau_t + (1-\lambda)(y_t^k - y_t^*) + u_t^k$$

	United States (1960-90)		OECD 1960-90		OECD 1990-2004	
OLS	0.48	(0.01)	0.93	(0.01)	0.75	(0.03)
Panel Dynamic OLS	0.47	(0.02)	0.91	(0.02)	0.74	(0.06)

Panel II: Fixed effect regressions

$$(c_t^k - c_t^*) = const + \tau_t + \phi^k + \beta_U(y_t^k - y_t^*) + u_t^k$$

	United States (1960-90)		OECD 1960-90		OECD 1990-2004	
OLS	0.50	(0.02)	0.98	(0.02)	0.63	(0.02)
Panel Dynamic OLS	0.52	(0.04)	0.99	(0.04)	0.66	(0.03)
Cointegration tests	-2.36		-2.42		-2.17	

NOTES: The results reported for the panel dynamic OLS estimation are based on estimating equations of the form $c_{kt} - c_{kt}^* = \widehat{b}x_{kt} + \sum_{l=-p}^{p} \delta_{kl}\Delta x_{t-l} + v_{kt}$ where $x_{kt} = (y_{kt} - y_{kt}^*)$ and $v_t^k = \tau_t + u_t^k$ or $v_t^k = \tau_t + \phi_k + u_t^k$, depending on whether it is a pooled or fixed regression. Standard errors are given in parentheses. Those for the PDOLS estimates are based on Mark and Sul (2003).All regressions control for time fixed effects. The panel cointegration tests at the bottom of the table are Pedroni's (20004) group mean t-statistics and are based on the PDOLS fixed-effect regressions.

Table 2: The increase in long-run risk sharing – robustness checks

Panel I: Different country groups and sample periods

Panel dynamic OLS estimates of $(c_t^k - c_t^*) = const + \tau_t + (1-\lambda)(y_t^k - y_t^*) + u_t^k$

Country group	1960-90		1975-90		1990-2004	
	pooled	FE	pooled	FE	pooled	FE
All	0.91	0.98	0.93	0.86	0.74	0.65
	39.38	27.27	25.91	10.96	12.29	22.35
All w/o most open SOEs	0.92	1.03	0.90	0.86	0.74	0.86
	(29.46)	(26.82)	(16.39)	(10.41)	(8.796)	(16.12)
SOEs only	0.89	1.04	0.90	1.06	0.73	0.65
	(22.54)	(11.62)	(17.26)	(10.00)	(12.32)	(29.47)

Panel II:long-horizon differenced regressions

OLS estimates of $(c_{t+l}^k - c_{t+l}^*) - (c_t^k - c_t^*) = const + \tau_t + \phi^k + \beta_U \left[(y_{t+l}^k - y_{t+l}^*) - (y_t^k - y_t^*)\right] + u_t^k$

Country group	1960-90		1975-90		1990-2004	
	3 yrs	5yrs	3 yrs	5yrs	3yrs	5yrs
All	0.83	0.84	0.90	0.78	0.78	0.79
	(14.65)	(11.27)	(10.73)	(7.93)	(10.74)	(10.61)
All but SOEs	0.88	0.95	0.92	0.88	0.79	0.78
	(14.30)	(11.56)	(12.74)	(12.92)	(17.22)	(10.09)
SOEs only	0.93	0.92	0.93	0.73	0.76	0.80
	(8.60)	(5.44)	(7.70)	(4.74)	(7.76)	(7.66)

Notes: Table provides estimates of the risk sharing coefficients based on differents sample periods, estimation methods and country groups. FE denotes the fixed effect estimate. 'All' refers to all 23 OECD countries, 'All but SOEs ' excludes the most open small economies Luxemburg, Ireland, Iceland and Switzerland. 'SOEs only' comprises 11 small open economies: Austria, Belgium, Denmark, Finland, Iceland, Ireland, the Netherlands, Norway, Sweden, Switzerland.

Table 3: Risk sharing patterns based on predicted and actual income data

Data	United States		OECD			
			1970-90		1990-2004	
	$1-\lambda = 0.48$		$1-\lambda = 0.90$		$1-\lambda = 0.74$	
	ex ante	*ex post*	*ex ante*	*ex post*	*ex ante*	*ex post*
	Panel I: Levels Regressions					
	β_K	β_C	β_K	β_C	β_K	β_C
predicted	0.45	0.03	0.08	-0.01	0.29	0.06
	(0.01)	(0.04)	(0.01)	(0.06)	(0.01)	(0.03)
actual	0.45	0.03	0.01	0.06	0.26	0.09
	(0.03)	(0.03)	(0.05)	(0.06)	(0.04)	(0.04)
	Panel II: Differenced Regressions					
	β_K^Δ	β_C^Δ	β_K^Δ	β_C^Δ	β_K^Δ	β_C^Δ
predicted	0.47	0.37	0.08	0.18	0.28	0.05
	(0.01)	(0.03)	(0.001)	(0.04)	(0.01)	(0.04)
actual	0.39	0.46	0.03	0.24	0.06	0.27
	(0.01)	(0.03)	(0.02)	(0.04)	(0.03)	(0.05)
	Panel III: Average Correlations between actual and predicted income growth					
	0.77		0.94		0.85	

NOTES: regression coefficients from equations (7) (*ex ante*) and (8) (ex post), based on actual and predicted income data. Panel I reports these regressions in level, panel II their first order differenced version. The predicted income data are generated according to equation (2) with the portfolio shares λ given at the top of the column. All regressions control for time-specific and region- or country specific fixed effects. Under the heading '*correlations*' we report the average across countries or regions of the correlation between predicted and actual idiosyncratic income growth.

Table 4: Long-term risk sharing and growth in international asset positions

Panel I: Regressions with Interactions

Coefficient on	1975-2004		1975-90		1990-2004	
	I	II	*I*	II	I	II
	no trend	trend	no trend	trend	no trend	trend
$\left[y_t^k - y_t^*\right]$	1.01	0.99	0.92	1.15	0.8779	0.99
	(19.97)		(10.67)	(6.25)	(17.37)	(4.98)
$\overline{GFA}_t \times \left[y_t^k - y_t^*\right]$	-0.1	-0.01	-0.02	-0.19	-0.05	-0.16
	(-2.94)		(-0.30)	(-0.85)	(-2.25)	(-2.70)
$trend \times \left[y_t^k - y_t^*\right]$		-0.01		0.01		0.04
				(0.87)		(2.77)

Panel II Regressions based on synthetic panels

	1975-2004		1975-90		1990-2004	
	low GFA	high GFA	low GFA	high GFA	low GFA	high GFA
$\left[y_t^k - y_t^*\right]$	1.004	0.86	1.09	0.97	0.93	0.75
	(30.15)	(33.91)	(31.42)	(38.42)	(17.47)	(18.48)

NOTES: Panel I reports results for PDOLS regressions of the form

$$c_{kt} - c_t^* = \left(1 - \lambda_t^{GFA}\right)\left[y_t^k - y_t^*\right] + \tau_t + u_{kt}$$

where $1 - \lambda_t^{GFA} = \kappa_0 + \kappa_1 \overline{GFA}_t + \kappa_2 t$ (column I) or $1 - \lambda_t^{GFA} = \kappa_0 + \kappa_1 \overline{GFA}_t$ (column II). Panel II reports OLS regressions of the form $c_t^k - c_t^* = (1 - \lambda)\left[y_t^k - y_t^*\right]$ for synthetic panels based on the highest / lowest quartile of the distribution of GFA_t^k as described in the main text, footnote 16, p. 32. t-statistics in parentheses.

Table 5: Risk sharing patterns as function of international asset positions

Panel I: Regressions with interactions

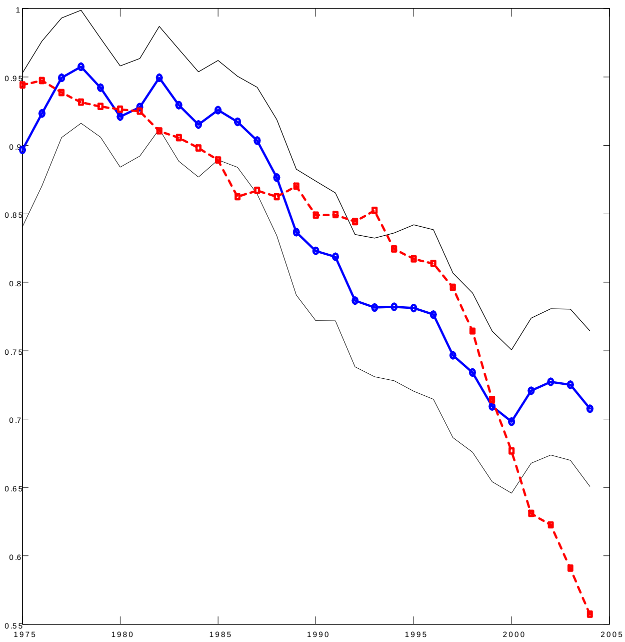
Interaction of $\left[y_t^k - y_t^*\right]$ with		total assets (GFA)		equity assets (GFE)		debt assets (GFD)	
		ex ante $\beta_K^k(t)$	*ex post* $\beta_C^k(t)$	*ex ante* $\beta_K^k(t)$	*ex post* $\beta_C^k(t)$	*ex ante* $\beta_K^k(t)$	*ex post* $\beta_C^k(t)$
1	(β_0)	**0.16**	**0.47**	**0.33**	**0.41**	*0.08*	**0.22**
		(4.80)	(8.30)	(16.28)	(8.94)	(1.95)	(3.31)
GFA_t^k	(β_{X1})	**0.08**	**-0.03**				
		(12.69)	(-3.05)				
GFE_t^k	(β_{X1})			**0.21**	**-0.16**		
				(26.73)	(-12.47)		
GFD_t^k	(β_{X1})					**0.09**	**0.05**
						(10.40)	(3.76)
trend	(β_{X2})	**-0.01**	0.01	**-0.02**	**0.01**	**-0.02**	0.0002
		(-4.32)	(0.25)	(-11.49)	(3.84)	(-11.07)	(0.08)

Panel II: Regressions based on synthetic panels

Panel group (sorted by GFx)	total assets (GFA)		equity assets (GFE)		debt assets (GFD)	
	ex ante $\beta_K^k(t)$	*ex post* $\beta_C^k(t)$	*ex ante* $\beta_K^k(t)$	*ex post* $\beta_C^k(t)$	*ex ante* $\beta_K^k(t)$	*ex post* $\beta_C^k(t)$
low GFx	**0.15**	0.03	**0.11**	0.11	**0.15**	0.04
	(3.59)	(0.48)	(3.19)	(1.76)	(3.05)	(0.54)
high GFx	**0.25**	**0.18**	**0.27**	**0.16**	**0.29**	**0.16**
	(7.22)	(6.80)	(7.13)	(4.96)	(7.23)	(4.90)

NOTES: Panel I reports results for PDOLS regressions of the form $\left[y_t^k - y_t^*\right] - \left[inc_t^k - inc_t^*\right] = \theta_k^K + \beta_{Kt}^k \left[y_t^k - y_t^*\right] + v_t^K$ and $\left[inc_t^k - inc_t^*\right] - \left[c_t^k - c_t^*\right] = \theta_k^C + \beta_C \left[y_t^k - y_t^*\right] + v_t^C$ where $\beta_{Xt}^k = \beta_{0X} + \beta_{1X} GFx_t^k + \beta_{2X} t$ for $X = K, C$ and where GFx stands for GFA, GFE, GFD in turn. Panel II reports OLS regressions of $\left[y_t^k - y_t^*\right] - \left[inc_t^k - inc_t^*\right]$ (for β_K) and $\left[inc_t^k - inc_t^*\right] - \left[c_t^k - c_t^*\right]$ (for β_C) on $\left[y_t^k - y_t^*\right]$. These regressions are based on synthetic panels formed from the highest and lowest quartile of the distribution of GFx at each point in time t according the the procedure described in the main text, footnote 16, p. 32. t-statistics in parentheses.

Figure 1: The increase in consumption risk sharing 1975-2004.



Notes: The blue (solid /dots) line is the sequence of cross-sectional estimates of $(1-\lambda_t)$. The red (dashed/ squares) line is $1-\lambda_t^{GFA} = 1.01 - 0.1\overline{GFA}_t$ where $\overline{GFA}_t$ is the cross-country mean gross foreign asset position. The thin (black) solid lines are the plus/minus two standard deviation bands for $1-\lambda_t$. These standard deviations are obtained using a jackknife resampling procedure.