Home Bias and the Structure of international and regional Business Cycles.

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Abstract

Consumption based tests suggest that industrialised countries share much less of their idiosyncratic risk than do U.S. federal states. Does this mean that capital markets are less integrated among countries or can this finding also be ascribed to different business cycle structures? Using the construct of a world portfolio of Shiller securities, we find that the income of US federal states is derived to about 50 percent from own output, that of OECD to about 80 percent. Against the benchmark of US federal states, the home bias in OECD country portfolio appears much lower than is commonly thought. The empirical regularity that consumption risk sharing appears much weaker in international data than in regional data sets can at least partly be ascribed to differences in the structure of international and regional business cycles.

Keywords: Consumption Risk Sharing, International and regional business cycles, Non-stationary panel data, structural VARs

JEL-Classification: F41, F43

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1 Introduction

Regressions of relative consumption growth rates on relative output growth rates suggest that capital markets are a lot more segmented between countries than between regions within a country.\(^1\) If countries or regions were completely diversified, the coefficient of such a regression, the ‘risk sharing coefficient’, should be zero. However, in country-level data, the risk sharing coefficient is typically found to be closer to unity than to zero. While it is generally a lot lower in regional data, the null of full risk sharing is typically rejected in both regional and international data sets (see e.g. Mace, Asdrubali, Sorensen and Yosha, Sorensen and Yosha, Hess and Shin (1998).

What the literature so far has failed to recognize is that under incomplete risk sharing, the magnitude of the risk sharing coefficient will depend on both a country’s or region’s portfolio structure and the structure of output fluctuations at home and in the rest of the world. Comparisons between the extent of risk sharing achieved at the regional or international levels will therefore have to account for differences in the structure of idiosyncratic business cycles.

In this paper, we ask to what extent the risk sharing coefficient is determined by the structure of business cycle risk and to what extent it is due to a generic home bias in country or region portfolios. The answer to our question may shed light on whether capital markets are really more complete between regions than between countries or whether differences in the structure of idiosyncratic risk can account for the apparent lack of international consumption risk sharing.

In our empirical analysis we build on Asrubali, Sorensen and Yosha (1996) and Crucini (1999) to distinguish between the degree of ‘diversification’ and the degree of ‘insurance’ that this diversification achieves. Diversification measures to what extent a country’s wealth is invested into a mutual fund of perpetual claims on world output (so-called Shiller (1993) securities). ‘Insurance’, defined as the magnitude of the ex-post correlation between relative consumption and output growth, measures to what extent the given portfolio structure actually helps to stabilise consumption across states of nature. In other words: diversification reflects the international portfolio structure whereas ‘insurance’ pertains to the income flows that this portfolio structure sets in motion, given the stochastic environment (i.e. regional or international business cycles).

Understanding the link between a macroeconomic notion of portfolio structure and the structure of international and interregional business cy-

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\(^1\)See Asdrubali, Sorensen and Yosha (1996) and Crucini (1999).
cles is paramount in understanding the sources of the observed differences in insurance at the regional and the national level. Certain types of risk may simply not be insurable, however, they may simply be less pervasive in, say, a regional data set than in an international one. One important dimension along which this may be the case is the extent to which idiosyncratic shocks are permanent.

There are good theoretical reasons to believe that permanent asymmetric fluctuations are hard to insure in existing financial markets. Because state-contingent contracts rely on the state of the world to be verifiable, they are generally harder to enforce than non-state contingent contracts. Kehoe and Perri (2002) have shown how enforceability constraints may give rise to endogenous market incompleteness. It is, however, not clear that these enforceability constraints (and the ensuing degree of market incompleteness) are a priori more binding between countries than between regions, say because of different legislatures or because of governments that may be biased towards taxing foreign capital owners. Rather, it is important to ascertain, whether differences in idiosyncratic persistence can potentially account for differences in the degree of insurance achieved by a country or region.

The distinction between diversification and insurance can be important in other contexts as well: Cole and Obstfeld (1991) study the case of a two-country world in which consumers have Cobb-Douglas utility. Each of the two countries specializes in the production of one differentiated good. In the Cole-Obstfeld example, there is complete insurance without any prior trade in assets because the real value of countries’ income is insured via terms-of-trade fluctuations.

The role of dynamic hedging mechanisms is emphasized in Lane (2001): the sale and purchase of state-contingent assets can provide insurance through capital gains. If countries or regions liquidate capital gains to finance current consumption, the may be able to obtain a high degree of of insurance with a huge (average) home bias in asset portfolios.

Our results suggest that the distinction between ‘diversification’ and ‘insurance’ is empirically important in understanding the order of magnitude of the home bias. US states are diversified to about 50 percent, whereas OECD countries have a degree of diversification of only 20 percent. These numbers indicate that there is still a home bias in international data, but that it is relatively smaller than is commonly thought. U.S. federal states are themselves far from being completely diversified. Our measures of the home

\textsuperscript{2}These are the possible explanations given in Kehoe and Perri

\textsuperscript{3}Idiosyncratic persistence may, in itself be the outcome of market incompleteness, in particular if prices are sticky. See Ghironi (2001/02).
bias are in line with interregional capital income flows in the United States. However, we find that the assumption of a constant international portfolio share is cannot be upheld in OECD data.

While our results indicate that the outcome of basic risk sharing regressions depends crucially on the structure of idiosyncratic risk, differences in the structure of international and (intra-) U.S. business cycles cannot account for the apparent lack of consumption risk sharing. Rather, it seems that OECD countries are much less successful than U.S. states in obtaining insurance against permanent idiosyncratic risk. This finding provides strong evidence in favour of the mechanism advocated in Keohoe and Perri: if international contracts have to be self-enforceable, permanent shocks may be uninsurable because they are more likely to make a country’s participation constraint binding than a transitory shock of equal size.

The remainder of this paper is structured as follows: in section two we discuss the measurement of risk sharing using consumption data. In section three we introduce the model and its implications for the measurement of risk sharing and diversification. In section four we present our data and discuss our results. Section five concludes.

2 Measuring risk sharing

In a world with complete capital markets, countries and regions will insure completely against any idiosyncratic risk. Therefore, ex post, there should not be any correlation between a country’s or region’s relative output and consumption. This fundamental insight was first applied to household level data by Mace (1991) and Cochrane (1991) and Townsend (1994). These authors suggested to run regressions of idiosyncratic consumption growth on idiosyncratic income growth.

In country-level data, household income is typically replaced by per capita GDP. The basic risk sharing regression to which we will refer throughout this paper will therefore have the form:

$$\Delta c_k^t - \Delta c_*^t = b \Delta y_k^t - \Delta y_*^t + \epsilon_t$$

where $c_k^t$ and $y_k^t$ are the logarithms of consumption and GDP in country $k$ respectively and $\epsilon_t$ is a disturbance term. ‘Rest of the World’-variables are denoted by an asterisk. Under the null of complete markets this regression should yield a coefficient of zero.

The acknowledgment that real world financial markets are likely to be incomplete in many ways has subsequently led to a more pragmatic approach in applied work. Rather than testing the null of complete markets, i.e. $b = 0$,
Asdrubali, Sorensen and Yosha (1996) as well as Sorensen and Yosha (1998) have argued very convincingly that the coefficient in the risk sharing regression may be of interest in itself and that it should be interpreted as a measure of the extent of risk sharing. Applying this insight to US state level data, ASY find that roughly 30 percent of idiosyncratic output fluctuations remain uninsured. Conversely, Sorensen and Yosha (1998) show that among OECD countries, more than 70 percent of idiosyncratic fluctuations remain uninsured.4

The basic risk sharing regression is motivated by a benchmark model with complete markets. The coefficient $b$ is then a natural metric of the deviation from the complete markets outcome. However, one has to be wary not to directly interpret $b$ as a measure of international diversification. In this paper, we introduce a distinction between the degree of diversification and the degree of insurance that a given degree of diversification may achieve. The reason for this distinction is that once we allow for imperfect international diversification, the structure of idiosyncratic fluctuations will matter for the extent of insurance that is eventually achieved. While $b$ measures the extent of insurance, we will show that it does not necessarily measure the degree of diversification. In other words: once markets are incomplete - for whatever reason - the allocation that can be obtained will depend on the stochastic environment, i.e. the structure of business cycles.

If we want to use the difference of country-level and regional measures of risk sharing to study the home bias, we will therefore need to gauge the extent to which different business cycle structure may account for the differential success in obtaining insurance.

Why do we not base our notion of the degree of diversification on a measure of net foreign assets? Firstly, data on cross-regional portfolio holdings do not exist, at least not on a broader basis and over sufficiently long time horizons. It is therefore impossible to gauge the extent of the national home bias relative to a (potentially present) regional home bias. Secondly, for our purposes, it may be problematic to classify a given asset as domestic or foreign. While this may be quite simple with respect to the equity of the local barber shop down the road (unless it is run by an international or interregional chain), it is clearly problematic with respect to multinational companies that may be listed on one country’s stock exchange but own subsidiaries or plants in many countries.

We therefore prefer to use the construct of a fictitious portfolio of Shiller-securities to measure a country’s or regions diversification. While such a

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4 The results in these two papers represent a flow-equivalent to French and Poterba’s (1991) observation of a huge home bias in international equity portfolio holdings.
portfolio may not have a directly observable counterpart in the real world, Shiller (1993) has shown very convincingly that it is a highly useful construction of the mind and we argue that contrasting this ‘portfolio’ with standard consumption based measures of risk sharing can shed light on the sources of the home bias and the structure of international and interregional risk sharing arrangements. We turn to a formal exposition of our model in the next section.

3 Accounting for business cycle structure

This subsection takes us to the core of the question asked in this paper: can differences in the structure of business cycles explain the different outcomes of the basic risk sharing regression? In order to address the problem, we will follow Crucini (1999) in motivating the risk sharing regression from a small theoretical model. While the focus in the earlier literature (including Crucini) is on regressions in relative growth rates, our focus here will be on panel regression of relative (logarithmic) consumption and output levels. This approach allows us to take account of the lower-frequency fluctuations of the data which are clearly important for the question we are addressing.

In presenting our approach, it will be useful to emphasize the distinction between ex-ante and ex-post risk sharing. International and interregional diversification pertains to an ex-ante sharing of risk through the cross-border holdings of assets. {ASY (1996) were the first to empirically explore this distinction.}

Ex post, that after the realization of this period’s uncertainty, consumption should be determined by permanent income. Under full risk sharing, country k’s consumption should correspond to world permanent income, which equals world permanent output:

\[ C_t^k = Inc_t^{kp} = \frac{r}{1 + r} \sum_{i=0}^{\infty} \frac{1}{(1 + r)^i} E[Y^*_t] = Y_t^{sp} \]

where \( r \) is the world real interest rate. Once we recognize that ex-ante risk sharing may be incomplete country k’s permanent income should be a weighted average of world and home permanent output:

\[ C_t^k = Inc_t^{kp} = \lambda Y_t^{sp} + (1 - \lambda)Y_t^{kp} \]

Here \( \lambda \) can be interpreted as the degree of diversification into perpetual claim on per capita output in the rest of the world. Using that world
consumption equals world permanent income, we can write

\[
\frac{C^k_t}{C^*_t} = \lambda + (1 - \lambda) \frac{Y_{t}^{kp}}{Y_{t}^{*p}}
\]  

(2)

This equation is reminiscent of Mace (1991), Asdrubali, Sorensen and Yosha (1996) or Cochrane (1991). It relates relative consumption to relative output. Under full risk sharing, the coefficient on relative (permanent) output should be zero. However, we note that this regression is in relative levels rather than relative growth rates. The second important difference is that the right hand side of (2) contains a non-observable variable, namely the relative permanent levels of home and foreign output. This feature of equation (2) will prove crucial for the problem we have at hand: our approach explicitly recognizes that whenever a country or region is not fully diversified, i.e. whenever \( \lambda < 1 \), the response of relative consumption to movements in relative output will depend on the structure of business cycles at home and in the rest of the world. In other words: while we can think of \( \lambda \) as the extent of portfolio diversification or of a measure of long-run insurance, the actual degree of insurance that a given degree of incomplete diversification will provide, depends on the structure of international macroeconomic fluctuations.

To illustrate this point, let us consider a simple example in which home and foreign output follow an AR(1) process:

\[
Y_t = \rho Y_{t-1} + \varepsilon_t \\
Y^*_t = \rho^* Y^*_{t-1} + \varepsilon_t
\]

where \( \rho \) and \( \rho^* \) are between zero and one. Using the definition of permanent income we get

\[
Y_{t}^{kp} = \frac{r}{1 + r} \sum_{l=0}^{\infty} \frac{1}{1 + r} \cdot E \left[ Y_{t+l}^{k} \right] = \frac{r}{1 + r} \sum_{l=0}^{\infty} \frac{1 + \rho^l}{1 + r} Y_{t}^{k} = \frac{r}{1 + r - \rho} Y_{t}^{k}
\]

In analogy we get

\[
Y_{t}^{sp} = \frac{r}{1 + r - \rho^*} Y_{t}^{*}
\]
So, the cake-sharing relation becomes

\[
\frac{C^k_t}{C^*_{t}} = \lambda_k + (1 - \lambda) \frac{1 + r - \rho^* Y^k_t}{1 + r - \rho Y^*_t}
\]

(3)

Hence, a cross-sectional or panel regression of the form

\[
\frac{C^k_t}{C^*_{t}} = \mu + \beta \frac{Y^k_t}{Y^*_t} + u_{kt}
\]

(4)

will not identify \((1 - \lambda)\) but rather we will have

\[
\beta = (1 - \lambda) \frac{1 + r - \rho^*}{1 + r - \rho}
\]

Certainly, \(\beta\) reflects how much of the idiosyncratic variance in output spills over into fluctuations in relative consumption. I.e. \(\beta\) measures the extent of insurance that is achieved, given a degree of international diversification, \(\lambda\) and the structure of the transitory but persistent part of relative output.

How sizeable the differences between the extent of insurance achieved and the extent of portfolio diversification can be, can also be gleaned from the present example: the estimate of \(\beta\) proves to be extremely sensitive to \(\rho^*\) and \(\rho_k\). Assume that \(\rho^* = 0.9\) and \(\rho_k = 0.95\). Then with a world interest rate of 2 percent, \(\beta\) will exceed \((1 - \lambda)\) by more than 70 percent! The opposite would be true, if RoW output was more volatile than home output.

To obtain a direct estimate of \(\lambda\), we will need to estimate the permanent components of home and foreign outputs. In order to do this, we will allow for a very general stochastic process to describe the joint dynamics of \(Y^k_P\) and \(Y^*_P\). We turn to this in the next section.

It seems that equation (2) would suggest a test as to whether \(\beta = (1 - \lambda)\), i.e. that the extent of insurance corresponds to the extent of diversification. Using current instead of permanent values in (4) would induce a bias in the estimation of \(\lambda\). Under the null that \(Y_t = Y^P_t\), i.e. output follows a random walk, we should have that \(\mu + \beta = 1\). This is a valid tests in the framework of our model but it may be of limited use in reality. The problem is that we have started from a model in which people consume the certainty equivalent of their future endowment streams. In reality, relative consumption should also depend on the relative risk with the income path \(\lambda Y^*_t + (1 - \lambda) Y^*_t\) as opposed to the diversified path \(Y^*_t\). Therefore, we should expect \(\frac{C^k_t}{C^*_{t}}\) to have a country-specific mean that does not only reflect the degree of diversification but also a country-specific risk premium.
Introducing a country-specific risk premium will induce a country-fixed effect in the logarithmic version of the Cake-Sharing equation. We use this logarithmic version because most economic time series are integrated of order one with normal residuals only after a logarithmic transformation.

We write the level risk sharing equation

\[
\frac{C_k}{C^*} = \lambda + (1 - \lambda) \frac{Y_k}{Y^*} + \pi_k
\]

where \(\pi_k\) is the risk premium. We rewrite

\[
\frac{C_k}{C^*} = 1 + \frac{C^K - C^*}{C^*}
\]

and use the approximation

\[
\log \frac{C_k}{C^*} \approx \frac{C^K - C^*}{C^*}
\]

Doing the same for relative permanent outputs, we get

\[
1 + \log \frac{C_k}{C^*} = \lambda - (1 - \lambda) 1 + \log \frac{Y_k}{Y^*} + \pi_k
\]

\[
\log \frac{C_k}{C^*} = (1 - \lambda_k) \log \frac{Y_k}{Y^*} + \pi_k
\]

where \(\pi_k\) can now be removed as a country-specific fixed effect. Having lower-case letters denote logarithms, we can write

\[
c_k^* - c^*_t = (1 - \lambda) \frac{\xi_k}{\xi^*} - \frac{\xi_k}{\xi^*} + \pi_k
\]

Under the assumption that home and aggregate output follow a random walk, first differencing would take us directly to the basic risk sharing regression (1). However, at least in the short-run, there is no guarantee that either \(y_t\) or \(y^*_t\) approximates their permanent values. Therefore, the conclusions form differences regression may be very different. In the next sections, we examine equation (6) empirically.

4 Econometric implementation

4.1 Data

We apply our approach two data sets: one for U.S. states and one for a group of 23 OECD countries. All data are annual.
The US-data are gross-state product data from the Bureau of Economic Analysis (BEA) and range from 1977-1999. Consumption data at the state level is not available. Most authors (Asdrubali, Sorensen and Yosha, Hess and Shin, DelNegro) have therefore used retail sales data and we follow this example. Specifically, our data are from DelNegro (2002) which ranges from 1960-1995. The sample range for the US-data we use in estimation is therefore 1977-1995. Country-level data are from the Penn World Table, release 5.6. and range from 1950 to 1992. The countries included in our estimation are:


Most of these countries are OECD countries and we will refer to them under this label. As regards the US, we follow the general practice in the US regional business cycle literature and include all states except Washington D.C.

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.

4.2 First results - differences vs. levels

For good econometric reasons, the empirical literature on consumption risk sharing has focused on regressions of growth rates rather than levels. Both idiosyncratic output and consumption are typically found to be integrated processes. The risk sharing regression in levels is likely to be spurious in time series unless there is a cointegrating restriction between the two variables. In Panel data, on the other hand, differencing removes any individual specific trend or mean in the variables.

The first line in table 1 provides the results of basic risk sharing regressions for both U.S. and international data. Roughly 75 percent of idiosyncratic output variability remains uninsured in country-level data, whereas only 15-30 percent of idiosyncratic variability spills over into consumption in U.S. state level data, results that are broadly in line with the findings obtained by Asdrubali Sornsen and Yoshia and Sorensen and Yoshia.

The picture changes quite dramatically once we consider the risk sharing regression in relative log-levels. The results are in the second line of table 1. The level regression can be thought of as capturing the long-run properties of the data. Equation (6) tells us to regress relative consumption levels on
relative \textit{permanent levels} of output. We construct various measures of $y^P$ and $y^*P$ below. However, to obtain a first impression of the data, a regression on plain levels may still be informative. While first differencing emphasizes the higher frequencies of the data, the level regressions should give us a picture of the long-run. We should therefore expect that using plain instead permanent values of relative output has less impact on the estimate of the coefficient than it has in differenced data (an impression that is confirmed below).

Because relative consumption and output levels are likely to be integrated variables, country-by-country regressions could potentially yield spurious results. However, this is not a problem once we exploit the panel dimension of the data set, even though the data may not be cointegrated (see Phillips and Moon (2000)). In the panel regression, we remove country-specific means.

In US data, still roughly 50 percent of output variability is found to be insured, down from 75 percent in the differenced regression. In international data, the change between the coefficient from the differenced and the level regression is found to be relatively small. About 85 percent of the permanent relative output variability remains uninsured. Hence, for most countries, relative consumption closely mimics the relative level of output, even in the long run.

In spite of the risk of spurious regression, the mean of the coefficients of country-by-country regressions is still-well-defined. Table 1 therefore also reports this value. For both countries as well as federal states, the mean estimate is slightly above the panel regression with removed fixed effects. Still, the results carry the same message: in particular for the US, there is considerably less risk sharing in the long-run than in the short-run. At the same time, we detect very little risk sharing between countries both in the short-run and in the long-run.

In figures 1a and 1b, we provide the cross-sectional plot of the panel of country and state data respectively. The optical impression confirms the regression results: in international data, relative consumption and output levels almost move one to one in the long-run. But also the U.S. data set displays a substantial comovement in relative output and consumption levels.

These first results suggest a very interesting pattern: The average U.S. state is nowhere close to completely diversified. The estimated degree of diversification is, certainly, much higher than in country-level data, but there is much less of a difference than the basic risk sharing regression would suggest. We further examine this finding by looking at appropriately constructed measures of permanent output in the next subsection.
4.3 Estimating portfolio shares

A crucial aspect in estimating international or interregional portfolio shares from an equation (6) is to obtain measures of the permanent component of \( Y^{kP} \) and \( Y^{*P} \). We compare two alternative specifications for the permanent components of home and foreign output. First, we consider a univariate AR(1) process in growth rates of home and foreign output.

\[
\begin{align*}
\Delta y^k_t &= \rho_k \Delta y^k_{t-1} + \nu^k_t \\
\Delta y^*_t &= \rho^* \Delta y^*_{t-1} + \nu^*_t
\end{align*}
\]  

(7)

This specification implicitly assumes that there are no spillovers between home and RoW output. We therefore also consider a VAR specification in output growth rates. In this specification, we also take into consideration that the maintained hypothesis in this paper is that aggregate consumption equals permanent income. If this is the case, then aggregate consumption should be a sufficient statistics for expected future levels of output. We therefore use the methodology first suggested by Campbell and Shiller (1989) and include consumption as an endogenous state in the VAR.

Now let

\[
x_t = \begin{bmatrix} y^k_t \\ y^*_t \\ c^k_t \\ c^*_t \end{bmatrix}
\]

denote the vector of endogenous variables. Then we estimate the VAR-model

\[
\Delta x_t = A \Delta x_{t-1} + \varepsilon_t
\]

We note that our estimation equation (6) involves the logarithms of the permanent values. However, we base our estimation on a VAR in logarithms. In the appendix we show that the permanent value \( Y_{kt}^P \) can be written

\[
Y_{kt}^P = Y_t \left( 1 + \frac{\Delta y^k_{t+1}}{(1 + r)^k} \right)
\]

We can then write
In the case of the VAR-process, we can use the Hansen-Sargent-prediction formula to get

\[
E_t \left( \sum_{l=1}^{\infty} \frac{\Delta y_{kt+l}^* - \Delta y_{kt+l}^*}{(1 + r)^k} \right) = h_0 \cdot A^{1-} \cdot \left( I - \frac{1}{1 + r} A \right)^{-1} \Delta x_t
\]

where \( h' = \begin{pmatrix} 1 & -1 & 0 & 0 \end{pmatrix} \). In the case of the AR(1)-process we get an expression that is analogous to (3). In our estimations, we impose \( r = 0.02 \).

In table 2 we provide the results of the risk sharing regression with relative permanent levels, based both on the univariate as well as on the VAR specification. Both specifications yield very similar results and also confirm the results from the simple level regression in table 1: The US shares 50 to 60 percent of its idiosyncratic long-run risk, whereas OECD countries share only 20 percent of it.

In the light of our theory, the regressions on relative permanent levels are directly informative about the degree of diversification of a group of countries or regions. The estimates in table 3 suggest that diversification in the sense of a portfolio of country- or region- Shiller securities is by far less complete in U.S. state level data than the basic sharing regression would suggest. There is certainly a huge home bias in international portfolios, 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.

Our results provide a fresh perspective on the home bias, and they seem highly plausible: the equity of small firms or companies is most likely not traded on international markets. Furthermore, our theoretical setup does not explicitly incorporate the role of labour. Claims to the labour share of national or regional outputs are equally not traded and our estimates should reflect this.

The central question of this paper is how a given degree of portfolio diversifications interacts with the structure of macroeconomic disturbances to generate the stylized facts from the basic risk sharing regression in differences. To understand this interaction, it is important to know whether the portfolio shares that we have estimated are in line with cross-border income flows. This helps us to gauge the plausibility of our estimated portfolio shares (that, as
we have argued, are unobservable) and it should give us further insights into the nature of the home bias in both national and international data.

In the next subsection we therefore compare the flow evidence on ex-post and ex-ante risk sharing with the evidence on stocks that we have generated in this and earlier sections of this paper. In this exercise and in the remaining part of the paper, we will build on the specification in which the permanent levels are estimated from the VAR. While the result are generally very similar, the VAR-specification has the added benefit that it nests the univariate one.

4.4 The channels of ex-ante and ex-post risk sharing

ASY (1996) were the first authors to explore empirically to what extent consumption smoothing is achieved through either ex-ante insurance or ex-post smoothing. While ASY report that more than 40 percent of relative income variability gets smoothed ex-ante, Sorensen and Yosha (1998) report that this channel is virtually inactive in international data. The contribution of ex-post smoothing, is, however, comparable in both international and national data. Can this evidence be reconciled with estimates of international or interregional portfolio shares of around 50 percent in regional and 20 percent in international data, as we have obtained them from our analysis?

In their work, ASY and Sorensen and Yosha associate the ex-ante channel of risk sharing with cross-border capital income flows, as measured by the difference between GDP and GNP (income). Ex ante-risk sharing can therefore be thought of as income smoothing. Ex-post risk sharing is then the smoothing of consumption through borrowing and lending in credit markets.

Our framework formalizes these notions of ex ante and ex post risk sharing. Income smoothing is achieved through the ex-ante diversification into world mutual fund of Shiller securities. Smoothed (or ‘pooled’) income is given by

\[ INC_t^k = \lambda Y_t^* + (1 - \lambda)Y_t^k \]  

(9)

However, it is important to note that our measure of income is based on a constructed portfolio of abstract assets. It is by no means clear that our national or regional income data coincide with the national accounts income (GNP) data that are used in the analysis of ASY ans SY. Therefore, the relative roles of ex-ante and ex-post risk sharing can differ between our approach and theirs and it is therefore interesting to compare the approaches.

Such a comparison raises an subtle but potentially important issue: the association of international income flows with ex-ante risk sharing and of savings behaviour with ex-post risk sharing is not a perfect one. Ex-ante risk
sharing may also be achieved through sale and purchase of foreign (state-contingent) assets. If the price of these assets is appropriately correlated with income uncertainty at home, capital gains could contribute to insure a country’s consumption against adverse income shocks. Conversely, temporarily high home income could be used to buy foreign assets when they are cheap. This mechanism is emphasized in Lane (2001). The problem now is that capital gains made from sales of such assets would not be recorded in national income (GNP). Still they would lead to international capital flows through the current account. Because these capital flows contribute towards smoothing consumption, given income, they would appear to be contributing to ex-post smoothing if the metric of ASY and SY is used. This is certainly correct in the sense that the sale of assets takes place after the resolution of uncertainty about income in the current period, but it neglects that a capital gain has also been reaped with the sale and that this capital gain is due to an ex-ante build-up of international portfolio holdings.

The gist of this reasoning is that the GDP-GNP differential completely reflects ex-ante risk sharing only to the extent that one assumes constant international portfolio holdings. Our framework allows us to assess whether this is a plausible assumption: if national accounts data and income data generated from (9) give us the same message concerning the relative roles of ex-ante and ex-post risk sharing, then this should be evidence that constant international portfolio holdings, (a constant home bias) are consistent with the data.

To assess this issue, we run regressions of the type suggested by ASY (1996):

\[
\Delta y^k_t - y^*_t - \Delta \ln c^k_t - \ln c^*_t = \beta_a \Delta y^k_t - y^*_t + u_t
\]

Here, \(\beta_a\) measures the extent of ex-ante risk sharing and ‘inc’ denotes the logarithm of income constructed according to (9).

In analogy, we also run

\[
\Delta \ln c^k_t - \ln c^*_t = \Delta \ln c^k_t - \ln c^*_t = \beta_p \Delta y^k_t - y^*_t + v_t
\]

and \(\beta_p\) now measures ex-post risk sharing.

ASY also consider a fiscal transfer channel. We cannot identify such a channel in our setup. However, we note that most fiscal transfers are not discretionary but based on rules or laws that have been set ex-ante. We would therefore argue that fiscal transfers provide mainly ex-ante insurance.

Table (3) contains our results and juxtaposes them to the findings obtained in ASY and SY. In U.S. data, we do indeed find that the amount of
ex-ante risk sharing detected based on the generated date set corresponds exactly the amount of smoothing achieved through both capital markets and fiscal transfers in the ASY setup. This is strong evidence that a constant home bias of around 40 to 50 percent in U.S. state level data is entirely compatible with the (flow) evidence from national accounts data. It is therefore, perfectly legitimate to associate capital income flows with ex-ante smoothing in US data.

The results differ, however, at the international level. Based on the artificial data set, all the risk sharing is achieved ex-ante, not ex-post. Conversely, the results in SY ascribe all risk sharing to international credit markets, i.e. to the ex-post channel. This suggests that a model that assumes constant value weighted international portfolio shares is not appropriate to describe the features of international national accounts income (GNP) data (as examined in SY (1998)). We interpret this as evidence that the mechanism put forward by Lane may be much more important in international data than in regional data sets. The finding is also consistent with the 'home bias and high turnover' puzzle first documented in Tesar and Werner (1995): Average international portfolio holdings are small, but also very volatile.

### 4.5 Insurance of permanent and transitory shocks

In the introduction we discussed why permanent shocks are likely to be harder to insure than transitory ones: permanent variation in idiosyncratic income can only be insured through state-contingent assets, which may not exist for all contingencies, or may at least be harder to enforce because the state of the world on which the asset is contingent may be imperfectly observable.

Our framework allows us to examine to which degree countries and regions obtain insurance against permanent and transitory changes in relative output. A potential source of the home bias detected in the baseline regression could be that country-specific fluctuations have a larger permanent component than do region specific business cycles. It would then not be clear a priori that financial markets are more complete between regions than between countries. We think that this is an important question to ask, in particular against the backdrop of our earlier finding that US federal states are not nearly as diversified as the baseline regression would suggest.

We therefore run separate regressions of idiosyncratic consumption growth rates on relative growth rates of the the permanent and transitory components of output respectively. The two regressions

\[
\Delta c - \Delta c^\ast = b_P \xi + \Delta y^P - \Delta y^P + \xi_t
\]

\[
\Delta c - \Delta c^\ast = b_T \Delta y^T - \Delta y^T + v_t
\]
then give us two separate measures of how consumption is insured against permanent \((b_P)\) and transitory \((b_T)\) fluctuations. The permanent components are constructed in the way described in the previous subsections and the change in the transitory part is then just \(\Delta y_t^T = \Delta y - \Delta y^P\).

In table (4) we present the results from this exercise. In U.S. data we find that 85-95 percent of permanent variability is insured whereas a similar number obtains for transitory fluctuations. In U.S. data we cannot find evidence that there is a qualitative difference between permanent and transitory shocks to output in as far as their degree of insurability is concerned. This result is in line with earlier findings by ASY (1996) who document that idiosyncratic persistence does not seem to have a big effect on the overall extent of insurance in U.S. data but that regions with more persistent idiosyncratic fluctuations rather tend to insure \textit{ex-ante}.

The picture changes quite substantially once we turn to the regression with international data. OECD consumption is less insured against both permanent and transitory shocks than in is U.S. consumption. But while the difference is not very big for the transitory variation, the coefficient on permanent output variation tells us that only 50 percent of permanent idiosyncratic output variability is insured at the international level. The difference between our estimate of \(b_P\) at the national and the international level is around 0.35 to 0.4. This is almost the difference between the coefficient estimates in the baseline regression. It seems that the home-bias detected in the baseline regression largely reflects a failure of countries to insure against permanent idiosyncratic risk. Our results provide strong support for the approach by Kehoe and Perri (2002): permanent shocks are harder to insure predominantly at the international level, which does indeed indicate that there are quantitatively important frictions in international financial markets that are absent in regional financial markets.

Still, while there is a lot less insurance at the international than at the regional level, it is important to note that the baseline regression reports estimates of the overall amount of risk sharing that are below both the degree of risk sharing that we find for either permanent or transitory shocks. This finding is true in both regional and international data. Again, the difference can be ascribed to the structure of business cycles: to the extent that permanent and transitory shocks are not completely insured, the overall extent of insurance that is achieved will also depend on the covariance structure of permanent and transitory fluctuations in output. To understand the anatomy of this result, we will write the coefficient of the baseline regression as a function of the coefficients on the transitory and permanent parts of output respectively:

Let the tilde denote relative growth rates of a variable, i.e. \(b = \Delta c - \Delta c^*\)
and $b = \Delta y - \Delta y^*$. Then the regression coefficient $b$ of the baseline risk sharing regression can be written as

\begin{align*}
  b &= \frac{\text{cov}(b, b)}{\text{var}(b)} = \frac{\text{cov}(b, b^p) + \text{cov}(b, b^T)}{\text{var}(b^p) + 2\text{cov}(b, b^p) + \text{var}(b^T)} \\
  &= b_p + b_T \frac{\text{var}(b_T)}{\text{var}(b^p)} 1 + \frac{2\text{cov}(b^p, b^T)}{\text{var}(b_p)} + \frac{\text{var}(b_T)}{\text{var}(b^p)} -1 \\
  &= [\alpha b_p + (1 - \alpha)b_T] \frac{\text{var}(b^p) + \text{var}(b^T)}{\text{var}(b)}
\end{align*}

(10a)

(10b)

(10c)

where

\begin{align*}
  b_p &= \frac{\text{cov}(b, b^p)}{\text{var}(b^p)} \\
  b_T &= \frac{\text{cov}(b^T, b^T)}{\text{var}(b^T)}
\end{align*}

are the regression coefficients of idiosyncratic consumption on idiosyncratic changes in the permanent and transitory component of output and the weight $\alpha$ is given by

\begin{align*}
  \alpha &= 1 + \frac{\text{var}(b_T)}{\text{var}(b^p)} -1
\end{align*}

If the covariance between changes in the permanent and transitory component of output is zero, then the coefficient of the baseline regression is a weighted average of the extent of insurance achieved for either permanent or transitory variation in relative outputs. There is, however, no reason to believe that the covariance term will generally vanish. Therefore, the overall extent of insurance that is measured by the baseline Mace-Cochrane-ASY regression is again a function of the structure of macroeconomic fluctuations, in this case the covariance structure of permanent and transitory components (or, in other words: the covariance structure of innovations in trend and cycle). If the term $\text{cov}(b^p, b^T)$ is positive, the baseline regression will detect a lower $b$, hence more insurance; if the covariance is negative, we see less insurance.
While equation (10) has been derived without any restrictions from theory, the intuition behind it can be understood in the framework of the permanent income hypothesis: Suppose a country enjoys a permanent positive idiosyncratic shock against which it was not or only partially insured ex-ante. Assume for simplicity that the country is however completely insured against transitory shocks. According to the relative variant of the PIH we have considered, relative consumption should instantaneously adjust to the new permanent level of relative output. If the adjustment to the permanent level is gradual, then current consumption should overshoot current relative output fluctuations, making consumption appear more volatile than output. But a gradual adjustment in output just means that an increase (decrease) in the permanent level decreases (increases) the transitory component. Hence, the change in the transitory and the change in the permanent component are negatively correlated - and this case, equation (10) would indeed predict that, ceteris paribus we find less insurance.

Conversely, if the permanent positive shock is also associated with a positive change in the transitory component, then this implies that current output changes will be larger than permanent changes. Because consumption will mainly adjust to the permanent part, it will react less strongly than current output changes, making consumption appear more insured in the basic risk sharing regression.

The evidence we find points strongly to a gradual adjustment of output to its permanent level: the coefficient from the basic risk sharing regression is in all cases bigger than either the one on permanent or the one on transitory changes in relative output alone.

5 Conclusion

In this paper we have examined the link between the structure of business cycles and consumption insurance. Consumption based tests suggest that there is much less risk sharing at the international than at the regional level. However, risk sharing, both in international as well as in regional data sets seems incomplete.

To the extent that risk sharing is incomplete, the structure of business cycles may matter for the degree of insurance that is eventually achieved, given the degree of international diversification. We have used the thought experiment of a portfolio of Shiller securities to distinguish between the extent of international diversification (as measured by the value weighted share of a world mutual fund in a country’s income) and the degree of insurance. We find that U.S. federal states are much less diversified than simple risk sharing.
regressions would suggest: their home bias is about 50 percentage points. Conversely, in country level data, we find a home bias of about 80 percent. This is much lower than evidence based on international equity portfolio holdings (French and Porterba (1991)) would suggest. While standard risk sharing regressions find a lot less consumption insurance in international data, our findings also suggest that a sizable component may be due to differences in the structure of international business cycles.

In U.S. data, a constant home bias of around 50 percent is also consistent with flow evidence on cross-regional capital income flows and the relative roles of *ex-ante* and *ex-post* consumption insurance as examined by Asdrubali, Sorensen and Yoshia. In international data, however, the assumption of constant international portfolio holdings does not seem to be compatible with data on cross-border capital income flows. This suggests that, while international portfolio holdings are relatively low, they are also likely to be very volatile, in line with Tesar and Werner’s (1995) finding of high turnover in international portfolio holdings.

Our method has also allowed us to investigate whether international macroeconomic fluctuations represent an altogether different type of risk than national business cycles. This could be the case to the extent that idiosyncratic output fluctuations have a bigger permanent component in international data than in regional data. There are good theoretical reasons to believe that permanent shocks are less insurable in existing financial markets than transitory fluctuations. Therefore, different degrees of persistence of asymmetric business cycles may be responsible for the home bias, without markets *per se* being less complete between countries than within. However, we find that such structural differences cannot account for the home bias. Rather, we document that U.S. federal states are much more successful than OECD countries in achieving insurance in particular against permanent idiosyncratic output shocks. The size of the home bias (as measured by basic risk sharing regressions) can almost exclusively be ascribed to this factor. This is strong evidence in support of the recent quantitative-theoretical work by Kehoe and Perri (2002), who argue that financial contracts may be incompletely enforceable among countries but not among regions.

There is a wealth of studies that have looked at risk sharing at either the U.S. state or at the international level. With only a few exceptions (Hess and Shin, DelNegro) most of them reach the conclusion that there is much more risk sharing within countries than among them. Most of them also find that consumption risk sharing is generally incomplete, be it at the international or national level. However, the exact numbers vary widely across studies. While this is partly due to different data sets, the specifications employed are also often quite different (compare e.g. the setup in Asdrubali, Sorensen
and Yosha (1996) to Crucini (1999). It is a byproduct of this paper to have shown to which extent apparently slight changes in specification can have an impact on the outcome of risk sharing regressions. Comparing and reinterpreting the outcomes from different specifications of the risk sharing regression offers a way to understand consumption risk is allocated, both at the the national as well as the international level.

References


### Table 1: Basic Risk Sharing and Cake Sharing Regressions

<table>
<thead>
<tr>
<th>Regression</th>
<th>United States</th>
<th>OECD</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta c - \Delta c^* = b(\Delta y - \Delta y^*)$</td>
<td>0.34 0.15</td>
<td>0.77 0.71</td>
</tr>
<tr>
<td>$c - c^* = b(y - y^*)$</td>
<td>0.61 0.50</td>
<td>0.92 0.87</td>
</tr>
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</table>

### Table 2: Estimates of the home bias - permanent level regressions

<table>
<thead>
<tr>
<th>Regression</th>
<th>United States</th>
<th>OECD</th>
</tr>
</thead>
<tbody>
<tr>
<td>$c - c^* = b(y^p - y^s^p)$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>AR(1)</td>
<td>0.50 0.42</td>
<td>0.83 0.79</td>
</tr>
<tr>
<td>VAR</td>
<td>0.63 0.52</td>
<td>0.86 0.80</td>
</tr>
</tbody>
</table>

### Table 3: Ex-ante and ex-post risk sharing in comparison.

<table>
<thead>
<tr>
<th>Regression</th>
<th>United States</th>
<th>OECD</th>
</tr>
</thead>
<tbody>
<tr>
<td>$inc$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>ex-ante</td>
<td>0.53 (0.02)</td>
<td>0.26 (0.02)</td>
</tr>
<tr>
<td>ex-post</td>
<td>0.31 (0.21)</td>
<td>-0.03 (0.18)</td>
</tr>
<tr>
<td>ASY (QJE96)</td>
<td>SY (JIE98)</td>
<td></td>
</tr>
<tr>
<td>$gnp$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>ex-ante</td>
<td>0.52</td>
<td>0.03</td>
</tr>
<tr>
<td>ex-post</td>
<td>0.30</td>
<td>0.25</td>
</tr>
</tbody>
</table>
Table 4: Insurance against permanent and transitory shocks

<table>
<thead>
<tr>
<th>Regression</th>
<th>United States</th>
<th>OECD</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>mean panel</td>
<td>mean panel</td>
</tr>
<tr>
<td>$\Delta c - \Delta c^* = b_P (\Delta y^P - \Delta y^*P)$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$b_P$</td>
<td>0.13 0.04</td>
<td>0.52 0.48</td>
</tr>
<tr>
<td></td>
<td>(0.25) (0.17)</td>
<td>(0.18) (0.20)</td>
</tr>
<tr>
<td>$\Delta c - \Delta c^* = b_T (\Delta y^T - \Delta y^*T)$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$b_T$</td>
<td>0.11 0.12</td>
<td>0.19 0.24</td>
</tr>
<tr>
<td></td>
<td>(0.40) (0.24)</td>
<td>(0.34) (0.30)</td>
</tr>
</tbody>
</table>
Figure 1: 'Cake Sharing' in International Data 1950-90

Figure 2: Long-term risk sharing in US state level data