Economics Department

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# International Risk-Sharing in the Short Run and in the Long Run*

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**Abstract**

Using a panel of 23 industrialised countries, the paper investigates how short-run and long-run income risks are shared and how the source of uncertainty matters for the way this risk gets insured. Surprisingly, short-term and long-term output risks are found to be equally well insured. Transitory shocks get smoothed almost completely whereas permanent shocks remain 80 percent uninsured. We find a somewhat more important role for international capital markets than earlier studies. Whereas our results tie in with some recent theoretical insights and are consistent with empirical findings on home bias in international portfolios, they raise the question why permanent shocks are so hard to insure internationally.

**Keywords**: International Consumption Risk Sharing, European Integration, Panel Data, Panel Vector Autoregressions

**JEL Classification**: F41, F43

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# 1 Introduction

Do industrialised countries use the same channels to insure against long-term and short-term income risks? Do they insure in different ways against different types of shocks? This paper aims to provide an answer to these questions.

The starting point of our analysis is the observation that most countries' consumption risks do not seem to be internationally diversified. French and Poterba (1991) were the first to hint at the huge home biases in international equity portfolios. This non-diversification puzzle has been cast into various formulations that are not only based on stocks of foreign assets but also on flow variables. Most notably, Backus, Kehoe and Kydland (1992) demonstrated that international consumption correlations are too low to be explained by models with perfect capital mobility and complete asset markets.

A complementary perspective on international non-diversification is provided in a series of papers by Asdrubali, Sørensen and Yosha (1996) and Sørensen and Yosha (1998). Asdrubali, Sørensen and Yosha (1996), henceforth ASY, suggested a simple decomposition of income risk that allows the investigator to distinguish between the cross-sectional and the intertemporal dimension of risk sharing. The cross-sectional dimension is reflected in the cross-border ownership of state-contingent assets such as equity or in fiscal transfer schemes. The intertemporal dimension is reflected in borrowing and lending, i.e. in the use of national or international credit markets.

In US state-level data ASY (1996) find that 39 percent of shocks to gross state product are smoothed by capital markets, 13 percent are smoothed by the federal government, and 23 percent are smoothed by credit markets. Conversely, Sørensen and Yosha (1998) find that EU and OECD countries achieve much less cross-sectional risk sharing than do US states. Mélitz and Zumer (1999) extended the ASY study by including further exogenous variables like regional size and the real interest rate. Their results by and large corroborate those of ASY.

The findings of the line of research surveyed here - and to which this paper aims to contribute - may have far-reaching implications for the prospective workings of the European Monetary Union. A plethora of studies documents that, in terms of Mundell's (1961) classical criteria, Europe is much less of an optimum currency area than are e.g. the United States. In particular, macroeconomic shocks are generally found to be much less symmetric in Euroland than among US states. Against this background, finding out to what degree capital markets can contribute to the insurance of aggregate income and consumption risk has become a question of paramount importance: if shocks are asymmetric, perhaps they can be smoothed through sufficient risk sharing. A monetary union that experiences asymmetric macroeconomic disturbances may not appear optimal when measured against the classical OCA criteria but it may provide a huge pool of risks that can be optimally insured - as long as the channels mentioned above are actually available and do get exploited.

Sørensen and Yosha (1998) conclude quite negatively in this respect: given that European countries do not seem to exploit risk sharing opportunities, EMU could entail high welfare costs in the absence of intensified fiscal transfer mechanisms. Mélitz and Zumer on the other hand conclude that the start of monetary union will promote the sharing of risks via market channels.

In this paper, we extend the method of ASY (1996) to a fully dynamic framework[1]. In so doing, we use recently developed methods for the estimation of panel vector autoregressions. Our method allows us to assess how income uncertainty at short and long forecast horizons is insured. It also allows us to investigate how different types of income uncertainty get insured. The most important distinction to be made along these lines is the one between permanent and transitory shocks to income. Insurance against permanent idiosyncratic shocks requires perpetual claims on some sort of income that is negatively correlated

[1]Athanasoulis and van Wincoop (2000) provide an analysis of risk-sharing at various time horizons but their model does not allow the identification of risk-sharing channels.

with a country's own income stream. Conversely, transitory fluctuations can be completely smoothed through borrowing and lending. *A priori*, we should therefore expect that permanent shocks get insured through different channels than transitory shocks and our econometric model allows us to disentangle these two types of shocks with minimal identifying restrictions.

Our results can be summarized as follows:

- Short-term and long-term risks are insured in the same ways. The forecast horizon does not matter for either the choice of insurance channel nor for the extent to which income risk is insured overall.

- Insurance against transitory shocks to income is generally much better than against permanent shocks and is achieved largely through credit markets, i.e. the intertemporal risk sharing channel. This result ties in with the theoretical findings by Baxter and Crucini (1995) who demonstrated that, as long as shocks are not too persistent, the full risk sharing allocations that pertain in models with complete asset markets can be very well approximated in models that only feature non-state contingent borrowing and lending. Whereas Baxter's and Crucini's work provides a theoretical rationalization of our results, a recent empirical study by Kraay, Loayza, Serven and Ventura (2000) has demonstrated that countries' international portfolios are largely held in the form of international credit rather than equity. This is the empirical corollary that may explain the importance of the credit channel for the sharing of transitory income risks.

- Earlier results in the literature suggested that capital markets provide only a minimal share of the total consumption insurance that is achieved between countries. Even though the role of capital markets for consumption insurance remains limited once we condition on the type of shock, their role seems much more respectable than would appear from our unconditional dynamic setup or from the results obtained in Sørensen and Yosha (1998).

- There is some evidence of insurance of permanent shocks through the intertemporal channel. The reason for this could be that a big share of a country's GDP cannot be traded on capital markets, e.g. because labour income is non-insurable. This may give rise to precautionary savings. Athanasoulis and Shiller (2000) have shown how the degree of market incompleteness affects the incidence of precautionary savings. Our results lend further empirical support to their view.

- Overall, roughly 60 percent of income variability in industrialised countries remains uninsured, most of it due to a failure to insure against permanent fluctuations in income.

Our results may also have important implications for further research into the sources of the home bias. It is generally found that national capital markets do much better in providing insurance to regions than do international capital markets in providing insurance to countries (compare for example the results in Asdrubali, Sorensen and Yosha (1996) and Sorensen and Yosha (1998)). Also, international asymmetries in output fluctuations are generally much more persistent than intranational ones (see for example Chamie et al. (1994)). At the same time, our results reveal that the failure of international capital markets to provide insurance is particularly due to a lack of insurance of 'permanent' income risks. An important question that future research should address is therefore why international capital markets do so badly in providing insurance against permanent shocks.

The remainder of the paper is organized as follows:

Section two outlines our dynamic econometric model of international risk sharing. In our empirical implementation, we rely on a panel vector autoregression that we implement using the method suggested by Holtz-Eakin, Newey and Rosen (1988). Our approach has the advantage that we only have to rely on cross-sectional asymptotics. Furthermore, we can identify permanent and transitory shocks to output with only minimal identifying assumptions by exploiting equilibrium relations between the data. Section three describes the data and our empirical results. We report the results of our analysis in section four and we offer conclusions in section five.

## 2 A dynamic model of risk sharing

In this section we propose a dynamic econometric model that enables us to analyse how income risk is shared over time. More specifically, our model allows us to identify the relative roles of intertemporal (i.e. borrowing and lending) and cross-sectional smoothing (i.e. insurance through international capital markets)[2].

The starting point of our analysis is the following decomposition of the variance of per capita GDP-growth:

$$\begin{aligned} var(\Delta gdp_t|\mathcal{I}_{t-1}) &= cov(\Delta gdp - \Delta gnp, \Delta gdp_t|\mathcal{I}_{t-1}) \\ &\quad + cov(\Delta gnp - \Delta c, \Delta gdp_t|\mathcal{I}_{t-1}) \\ &\quad + cov(\Delta c, \Delta gdp_t|\mathcal{I}_{t-1}) \end{aligned} \tag{1}$$

Lower case letters denote logarithms and *gnp* and *c* denote gross national product and consumption per capita respectively. The conditioning information set $\mathcal{I}_{t-1}$ contains realizations of variables that are known at the end of period $t-1$.

We can divide (1) by $var(\Delta gdp_t|\mathcal{I}_{t-1})$ to get:

$$1 = \beta_C + \beta_S + \beta_U$$

where

$$\boldsymbol{\beta} = \begin{bmatrix} \beta_C \\ \beta_S \\ \beta_U \end{bmatrix} = \frac{1}{var(\Delta gdp|\mathcal{I}_{t-1})} \begin{bmatrix} cov(\Delta gdp - \Delta gnp, \Delta gdp) & \\ cov(\Delta gnp - \Delta c, \Delta gdp) & |\mathcal{I}_{t-1} \\ cov(\Delta c, \Delta gdp) & \end{bmatrix} \tag{2}$$

[2]Our method is closely related to Asdrubali, Sørensen and Yosha (1996). Their approach is completely static, however. Another difference between our model and that of ASY is that we do not allow for a fiscal insurance channel. Sørensen and Yosha (1998) have demonstrated that the fiscal channel is not important for the international dimension of risk sharing which is what we focus on in this paper.

The individual coefficients in $\boldsymbol{\beta}$ can now be associated with various channels of risk sharing. The GDP-GNP differential reflects international factor income flows. Hence, $\beta_c$ measures to what extent capital income from abroad covaries with GDP. Therefore, $\beta_c$ can be thought of as representing the cross-sectional dimension of consumption insurance that is achieved (primarily) through cross-border ownership of foreign assets, i.e. through international capital markets.[3]

The GNP-C differential measures savings and $\beta_s$ gives the contribution of the intertemporal aspect of consumption insurance (i.e. smoothing through savings). Finally, $\beta_U$ is the residual covariance between consumption growth and GDP growth, reflecting the undiversified or unsmoothed component of consumption.

We will now describe how we identify the conditional variances and covariances involved in (1). For this purpose, let

$$\Delta \mathbf{X}_t = \begin{bmatrix} \Delta gdp_t \\ \Delta gdp_t - \Delta gnp_t \\ \Delta gnp_t - \Delta c_t \end{bmatrix}$$

Then we assume that

$$\mathcal{I}_t = \{\mathbf{X}_\tau\}_{\tau=1}^{t} \tag{3}$$

and that expectations coincide with linear projections. These assumptions allow us to express $\mathbf{E}(\Delta X_t | \mathcal{I}_{t-1})$ as a vector autoregression. The unexpected component of $\boldsymbol{\Delta}\mathbf{X}_t$ which we will denote by $\varepsilon_t$ is now given by the reduced-form residual of the VAR:

$$\boldsymbol{\Phi}(\mathbf{L})\boldsymbol{\Delta}\mathbf{X}_t = \varepsilon_t \tag{4}$$

where $\boldsymbol{\Phi}(\mathbf{L})$ is a $3 \times 3$ matrix polynomial in the lag operator, $\mathbf{L}$, which satisfies the condition that the roots of $\det(\boldsymbol{\Phi}(\mathbf{z}))$ lie outside the unit circle.

[3] As Sørensen and Yosha (1998) note, labour income flows between industrialised countries are negligible. The same holds true for interest payments on international bonds and loans. We can therefore think of the GDP-GNP differential as a good proxy for contingent capital income such as equity returns.

Now let $\mathbf{\Omega}$ denote the variance-covariance matrix of $\varepsilon_t$ and let $\omega_{ij}$ be the entry in the $i$-th row and $j$-th column of $\mathbf{\Omega}$. Then

$$\beta_c = \frac{\omega_{21}}{\omega_{11}} \text{ and } \beta_S = \frac{\omega_{31}}{\omega_{11}}$$

Of course, the analogue of $\beta_u$ is given by

$$\beta_u = 1 - \beta_c - \beta_s \tag{5}$$

We can now generalize our approach to arbitrary forecast horizons in order to answer the question as to what the role of various channels for risk sharing at these horizons may be. The mean squared prediction error in a VAR, $k$ periods ahead, is given by

$$\mathbf{\Psi}(k) = MSPE_k = \sum_{l=0}^{k-1} \mathbf{C}_l \mathbf{\Omega} \mathbf{C}_l' \tag{6}$$

where the $\mathbf{C}_l$ are the matrix coefficients of the moving average representation of $\mathbf{\Delta X}_t$.

Let the entries of $\mathbf{\Psi}(k)$ be denoted by $\psi_{ij}(k)$. Then the analogue of $\boldsymbol{\beta}$ from above can be defined:

$$\beta_c(k) = \frac{\psi_{21}(k)}{\psi_{11}(k)} \text{ and } \beta_s(k) = \frac{\psi_{31}(k)}{\psi_{11}(k)}$$

and again

$$\beta_u(k) = 1 - \beta_c(k) - \beta_s(k)$$

and

$$\boldsymbol{\beta}(k) = \left[ \begin{array}{ccc} \beta_c(k) & \beta_s(k) & \beta_u(k) \end{array} \right]'$$

Obviously, $\boldsymbol{\beta}(1) = \boldsymbol{\beta}$ because $\mathbf{C}_0 = \mathbf{I}$ and therefore $\mathbf{\Psi}(1) = \mathbf{\Omega}$.

Note also that as the forecast horizon gets infinite, $\boldsymbol{\beta}(k)$ should converge to the unconditional $\boldsymbol{\beta}$ that emerges from the static ASY model. Hence, the basic ASY regression provides a check of specification for any VAR estimation that may provide the basis for the dynamic decomposition given in (6). We are now going to deal with estimation issues.

## 2.1 PVAR Estimation

A naive application to an individual country of the procedure outlined in the previous section, is not likely to yield meaningful results. Estimating separate VARs for each country would not allow us to control for country fixed effects, possibly leading to seriously biased estimates. Also, we need to take into account time varying fixed effects that are common to a whole cross-section of countries. This is because common or global shocks cannot be insured and we have to make sure that they do not pollute our estimates. We will therefore employ panel techniques in order to identify country-specific and time-specific components by exploiting not only the time series but also the cross-sectional dimension. As our sample period is relatively short - we employ annual data from 1975 to 1990 - we will rely on dynamic panel methods that are robust to short samples, i.e. require only cross-sectional asymptotics.

To see the problems that are associated with estimating this type of dynamic panel model, write out the standard reduced-form representation (4) to get

$$\mathbf{\Delta X}_t = \mu + \sum_{l=1}^{p} \Phi_l \mathbf{\Delta X}_{t-l} + \varepsilon_t \qquad t = p+1, ..., T \tag{7}$$

Reinterpreting this as a system of panel equations yields

$$\mathbf{\Delta X}_{it} = \mu + \sum_{l=1}^{p} \Phi_l \mathbf{\Delta X}_{i,t-l} + \lambda_t + f_i + \mathbf{u}_{it} \qquad i = 1, ..., K; t = p+1, ..., T \tag{8}$$

where now all variables vary by $i$ and $t$, and where $f_i$ is the vector of country-specific effects and $\lambda_t$ is a time-specific effect. Since $\mathbf{\Delta X}_{it}$ is a function of $f_i$, $\mathbf{\Delta X}_{i,t-1}$is also a function of $f_i$. Therefore, $\mathbf{\Delta X}_{i,t-1}$, a right-hand regressor in (8), is correlated with the error term. This renders the OLS estimator biased and inconsistent even if the $\mathbf{u}_{it}$ are not serially correlated. For the standard fixed effects (FE) estimator, the 'within' transformation wipes out the country-specific effects $f_i$, but $(\mathbf{\Delta X}_{i,t-1} - \bar{\mathbf{\Delta X}}_{i,.-1})$ where $\bar{\mathbf{\Delta X}}_{i,.-1} = \sum_{t=2}^{T} \mathbf{\Delta X}_{i,t-1}/(T-1)$ will still be correlated with $(\mathbf{u}_{i,t} - \bar{\mathbf{u}}_{i,.})$ even if the $\mathbf{u}_{it}$ are not serially correlated.

This is because $\bar{\mathbf{u}}_{i,.}$ contains $\mathbf{u}_{i,t-1}$ which is correlated with $\mathbf{\Delta X}_{i,t-1}$ by construction. In the technical appendix, we describe how we have used instrumental variables techniques following the method set out by Holtz-Eakin, Newey and Rosen (1988) to estimate the model given in (8).

In the remainder of this section, we are going to discuss how we can incorporate permanent and transitory shocks in the VAR-model (8). Since in what follows, panel notation will generally not be required, we will henceforth drop the index $i$ or the fixed effects in our discussion.

## 2.2 Permanent shocks and risk sharing

Our interest in this paper is in the comovement of *growth rates* of consumption and various output aggregates. Still,it is possible that the *levels* of these variables may have feedback effects on the growth rates. To the extent that output and consumption are likely to follow integrated processes, such feedbacks from the level-variables would imply cointegrating relations between the variables.

In standard models of economic growth cointegrating relationships will most likely arise in the form of a stationary 'great ratio' of consumption over output (see for example King et al. (1991) and Neusser (1991)). In our setup, the great ratio is just given by the difference $c - gnp$[4]. A formal test based on the dynamic panel OLS procedure suggested by Nelson and Sul (2001) strongly rejected the null of no cointegration between idiosyncratic GNP and consumption. We therefore decided to add this variable as an error correction term to the panel VAR model given in (14). Our model accordingly looks as follows:

$$\Delta X_{it} = \mathbf{\Phi}\Delta X_{it-1} + \boldsymbol{\gamma}\boldsymbol{\delta}' X_{it-1} + \varepsilon_{it} \tag{9}$$

where $\boldsymbol{\delta}' = \begin{bmatrix} 0 & 0 & -1 \end{bmatrix}$ is the cointegrating vector and $\boldsymbol{\gamma} = \begin{bmatrix} \gamma_1 & \gamma_2 & \gamma_3 \end{bmatrix}'$ represents the vector of adjustment coefficients.

[4]We measure the great ratio as $C/GNP$, not, as is common, as $C/GDP$. The reason for this is, that at least in principle, a country's GDP and consumption can diverge arbitrarily if foreigners own perpetual claims on a sufficiently large share of that country's income. This is exactly what we should see if risk sharing was perfect.

An error-correction model such as (9) allows the identification of permanent and transitory disturbances without further identifying assumptions. Following Johansen (1995), the permanent shocks can be written as

$$\pi_{it} = \gamma'_{\perp} \varepsilon_{it} \tag{10}$$

whereas the transitory disturbances are identified by requiring that they be orthogonal to the space of permanent shocks. Hence

$$\tau_{it} = \gamma' \Omega^{-1} \varepsilon_{it} \tag{11}$$

Note that whenever $\pi_t$ or $\tau_t$ are non-scalar, the permanent or transitory shocks are not identified among themselves. However, for our purposes, this does not matter. The share of the forecast error variance that is explained by all permanent or transitory shocks does not depend on how we identify each of these shocks individually.

To see this assume that $\mathbf{S}_\pi$ and $\mathbf{S}_\tau$ are appropriately dimensioned non-singular matrices such that $\pi_0 = \mathbf{S}_\pi \pi$ and $\tau_0 = \mathbf{S}_\tau \tau$. Let furthermore, as in the non-cointegrated case, $\mathbf{C(L)}$ be the reduced form matrix polynomial giving the Wold-representation of (9). Then the structural, i.e. just-identified form is given by $\mathbf{C(L)P}^{-1}$ where

$$\mathbf{P} = \begin{bmatrix} \mathbf{S}_\pi \gamma'_\perp \\ \mathbf{S}_\tau \gamma' \Omega^{-1} \end{bmatrix}$$

is just the matrix mapping the reduced-form disturbances into their permanent and transitory components.

It is easily verified that

$$\mathbf{P}^{-1} = \begin{bmatrix} \Omega \gamma_\perp (\gamma'_\perp \Omega \gamma_\perp)^{-1} \mathbf{S}_\pi^{-1} & \gamma (\gamma' \Omega^{-1} \gamma)^{-1} \mathbf{S}_\tau^{-1} \end{bmatrix}$$

Then note that the covariance of $\begin{bmatrix} \pi_0 & \tau_0 \end{bmatrix}'$ is given by

$$\mathbf{\Sigma} = \begin{bmatrix} \mathbf{S}_\pi \gamma'_\perp \Omega \gamma_\perp \mathbf{S}'_\pi & \mathbf{0} \\ \mathbf{0} & \mathbf{S}_\tau \gamma' \Omega^{-1} \gamma \mathbf{S}_\tau \end{bmatrix}$$

Hence, the mean-square prediction error is

$$\Psi(k) = \sum_{l=0}^{k-1} \mathbf{C}_l \left[ \boldsymbol{\Omega}\gamma_\perp(\gamma'_\perp\boldsymbol{\Omega}\gamma_\perp)^{-1}\gamma'_\perp\boldsymbol{\Omega} + \gamma(\gamma'\boldsymbol{\Omega}^{-1}\gamma)^{-1}\gamma' \right] \mathbf{C}'_l \qquad (12)$$

where the first term in parentheses measures the contribution of permanent shocks and the second the transitory. It can be seen from (12) that $\Psi(k)$ is independent of any particular choice of $\mathbf{S}_\pi$ and $\mathbf{S}_\tau$. Hence, the relative contributions of permanent and transitory shocks do not depend on the particular just-identification chosen.

We are going to report the estimation results for the cointegrated panel VAR and the ensuing decomposition of the prediction error in section 4.

# 3 Data and Empirical Implementation

We used annual per capita data for GDP, GNP and consumption (C), for 23 industrial countries, from the Penn World Tables (PWT, release 5.6). We generated world per capita aggregates of each of the three variables using population data from the same source. Annual observations on all three variables were available for the period 1970-90. In our estimation, we included only the period 1975-90 in order to avoid potential parameter instability in the model that is bound to arise if the oil shock and the aftermath of the demise of Bretton-Woods was included. Following Sørensen and Yosha (1998), we did not extend the sample beyond 1990 to avoid instability problems that are likely to arise from German unification. These limitations make the sample rather short, but our econometric methods, in particular the instrumentalisation following Holtz-Eakin, Newey and Rosen (1988), are designed to cope with small time dimensions.

We transformed all data into log first differences to generate growth rates. Then, to account for the potential role of global shocks that may create uninsurable output variability, we formulated the data for each country relative to the global aggregate. In the setup of the panel, we multiplied the data of each country by its population weight. The description of variables from the PWT data base and the list of countries are given in the data appendix.

We then proceeded to the panel estimation of the vector autoregressions given in (9), using the method suggested by Holtz-Eakin, Newey and Rosen, as described in the previous subsection and the technical appendix. In the estimation, we included time-specific fixed effects to account for any remaining cross-sectional dependence and individual-specific fixed effects.

We used standard information criteria to determine the lag length of the VAR-model. Those generally suggested 1-2 lags. As an additional test of specification, we used the fact that, as $k$ tends to infinity, $\boldsymbol{\beta}(k)$ should converge to the unconditional $\boldsymbol{\beta}$, i.e. the vector of coefficients of the simple panel regressions on $\Delta gdp$ of $\Delta gdp - \Delta gnp$, $\Delta gnp - \Delta c$ and $\Delta c$ respectively. This test generally required us to impose two lags which we used throughout.

We then inverted the VAR to generate forecast errors according to (6). We now discuss the results of this exercise.

# 4 Results

In the selection of countries we used for our investigation, we deliberately only included industrialised economies. This is to ensure that countries are sufficiently homogenous to warrant treatment in a single panel estimation. Our panel also includes several interesting sub-groupings and we will report results for these throughout. These sub-groups include the G7, the EU 15 and a core group of European economies. Again, the appendix provides more detail.

Table 1 provides the relative contributions of the intertemporal and the cross-sectional channels at various forecast horizons.

It is a first interesting feature of our results that the relative contribution of the channels does not vary over time. To save space, table

1 reports, the results for the one and three year horizons only but the findings at other horizons are virtually identical. This is a remarkable result that we found to be extremely robust to changes in the model specification. It may seem surprising that short-term and long-term risks are equally well insured. Intuitively, one might expect that various forms of capital market imperfections would lead to a 'short-term bias' in consumption insurance. This does not seem to be the case.

Our results may also suggest that a fully specified dynamic econometric model such as the one put forward in this paper is required to get at the issue of dynamic risk sharing. Earlier contributions to the literature which admittedly were not primarily concerned with the dynamics of risk sharing, tended to use a short-cut to gauge risk sharing at different horizons: they typically look at data differenced at high and low frequencies. Using this method, Sørensen and Yosha (1998) find that the unsmoothed component at the three-years horizon is much larger (roughly 75 percent) than at the 1-year horizon (roughly 60 percent). In a similar way, Canova and Ravn (1996) report that lower frequency fluctuations in income are less insured than higher frequency fluctuations. The results in our paper, including those to be reported below for permanent and transitory shocks, demonstrate that these exercises provide a good estimate of how well-insured countries are against shocks of various degrees of persistence but not how well-insured they are against risks at different horizons.

The last column of table 1 reports the estimate of the unconditional model, i.e. practically a re-run of the Sørensen and Yosha (1998) procedure on our data. These unconditional estimates display the same pattern that was already found by Sørensen and Yosha (1998). Capital markets virtually do not matter for risk sharing, the bulk of insurance is provided through (intertemporal) self-insurance. Interestingly enough, our conditional estimates from the dynamic model find a somewhat more important role for international capital markets. However, once one takes account of the estimation uncertainty in the unconditional model, the respective results are not too far apart.

Overall, we find that the conditional estimates eventually converge to unconditional ones - at least after taking account of the relatively large estimation uncertainty in the unconditional model. This is reassuring as it provides a check of specification of our dynamic model as has been suggested in section two.

The VAR-based approach we have suggested in this paper allows us to examine an important assumption that underlies the ASY-approach: if the GDP-GNP differential and the GNP-Consumption differential actually serve as buffers for shocks to output, they should be driven by exactly the same shocks that drive GDP. In other words; the notion underlying ASY and the related literature is that shocks originate in output fluctuations and get smoothed at various levels. But the various aggregates, i.e. the GDP-GNP differential and the GNP-C differential that act as buffers, should not themselves be the source of shocks.

We can examine this assumption by conducting a principal components analysis of the shocks to our econometric model. If the presumption underlying the ASY approach is correct, then there should be a single dominant principal component in the reduced form errors that we get from the estimation of the VAR. Furthermore, this principal component should be highly correlated with innovations in the $\Delta gdp$-equation of our model but virtually uncorrelated with innovations in the other two equations.

In table 2, we give the share of the total variation in $[\Delta gdp, \Delta gdp - \Delta gnp, \Delta gnp - \Delta c]$ that is explained by the first principal component of $\mathbf{\Omega}$. As it turns out, we *do* find a dominant principal component in the reduced-form errors for all groupings of countries that we examine. We then also calculated the correlation of this principal component with unexpected innovations in the $\Delta gdp$-equation, i.e. $\varepsilon_{1ti}$ as well as the $\Delta gdp - \Delta gnp$- and $\Delta gnp - \Delta c$-equations, $\varepsilon_{i2t}$ and $\varepsilon_{i3t}$ respectively.. These correlations are given in columns 2-4 of table 2. Our results suggest that, indeed, shocks to $\Delta gdp$ drive the joint dynamics of $[\Delta gdp, \Delta gdp - \Delta gnp, \Delta gnp - \Delta c]$. This is a very important finding as it demonstrates the validity of our method and the static versions of it that have been used in ASY (1996), Sørensen and Yosha (1998) and Mélitz and Zumer (1999).

## 4.1 Permanent and transitory shocks

In table 3 we provide forecast error decompositions for the different sources of income uncertainty, i.e. permanent and transitory shocks.

These decompositions are similar to the unconditional dynamic results we reported in table 1 in that they do not vary over time. However, our results also reveal that there are important differences in the way that the various channels contribute to the sharing of risks that arise from different sources of shocks.

Firstly, permanent shocks are insured to a much lesser extent than transitory shocks. This finding is in line with earlier results in Canova and Ravn (1996) who also found that low-frequency risks seem to be insured less than high frequency fluctuations. In fact, when the panel VECM is estimated with all countries included, transitory shocks are found to be almost perfectly smoothed. We note that, very much as in the unconditional case, the forecast horizon does not matter for the extent of total insurance nor for the relative role of the channels.

Secondly, once we consider the channels by which these shocks get insured, we find that insurance against transitory fluctuations is almost exclusively achieved through the intertemporal channel, whereas, in line with the findings by Sørensen and Yosha (1998), the role of capital markets remains limited.

The results for transitory shocks tie in with recent empirical research by Kraay, Loayza, Serven and Ventura (2000) that suggests that the international component of most countries' portfolios is heavily biased towards loans and bonds. On the theoretical side, Baxter and Crucini (1995) have shown that the full risk sharing allocations that ensue as equilibria in models with complete markets can be approximated by models that feature only non-state-contingent assets. This result holds as long as shocks are not too persistent. Our results highlight the empirical relevance of the Baxter and Crucini study: even though individual countries' international portfolios show a huge home bias, they seem sufficiently diversified to achieve almost full insurance against transitory output risks.

This insurance seems to be achieved largely with bonds and international credit rather than equity or other state contingent assets.

The results we obtained for permanent shocks are particularly interesting in two respects. First, when the model is estimated with data from all countries, the intertemporal and the cross-sectional channels play almost equal roles. Certainly, the bulk of permanent income risk remains uninsured, but the relative contributions to the amount of insurance that is eventually achieved is roughly equal for the cross-sectional and intertemporal channels. In particular, it is noteworthy that the intertemporal channel matters at all for the insurance against permanent risks. Models in which only the expected path of income matters for the savings decision will not be able to rationalize this feature of the data. Rather, income variability appears to matter in this case. Athanasoulis and Shiller (2000) have shown how the extent of observed precautionary saving depends on the degree of incompleteness of markets for claims on national income. Accordingly, we interpret our finding as evidence of precautionary savings.

When the model is estimated with only a subgroup of countries, our results are generally confirmed. One particular point may be worth mentioning, though:

The role of the cross-sectional channel, i.e. international capital markets, for the insurance against permanent shocks is less pronounced in all of the sub-groups than it is when the model is estimated with all 23 countries. The sub-groups are more homogenous in terms of country-size than is the whole panel. Our results could suggest that risk sharing through international equity markets is more pronounced between countries of different size. In this respect, our results are in line with Lane (2000) who has found that smaller countries tend to hold more foreign equity than do larger countries.

Summarizing this section, we can say that, in annual data, permanent output fluctuations account for just below eighty percent of total output variability and that only twenty percent of these fluctuations are insured. Hence, we find that at least 60 percent of total output variability is uninsured. This is in line with the results obtained by Sørensen and

Yosha. Our findings complement theirs in that it seems that most of this uninsured component is due to uninsured permanent shocks. This raises the question why international capital and credit markets do so poorly in insuring people's consumption against permanent shocks in income.

# 5 Conclusion

In this paper, we have investigated in which way industrial countries insure against output fluctuations. In so doing, we have offered two important novelties.

The first is that we consider how risks are insured at various horizons, thus providing a dynamic version of the method first proposed by Asdrubali, Sørensen and Yosha (1996). Our results corroborate the notion of a home bias in international risk sharing, even for forecast horizons as low as a year.

Secondly, we also find evidence that an important share of the variation in idiosyncratic output and consumption may be of a permanent nature. Permanent shocks need to be insured through perpetual claims. The French-Poterba observation of a home bias in international equity portfolios may suggest that most countries are badly insured against permanent fluctuations in their income streams. Once we allow for non-stationarities in our data set, our findings are consistent with this view: there is generally little insurance against permanent shocks but transitory risk in output is almost completely insured, mainly via national and international credit markets.

The second finding is in line with recent empirical evidence that suggests that international portfolios do not only display a home bias but are also severely biased towards non-state contingent assets such as bonds and loans. A theoretical rationalization for our results may be given by Baxter and Crucini (1995), who demonstrated that full risk-sharing allocations can be approximated quite well in economies with imperfect capital markets as long as shocks are not too persistent.

Our aim in this paper was to draw a map of an area of our ignorance, i.e. how countries share risks at various time horizons. We have not put forward any particular theory of what the intertemporal pattern of risk sharing should look like. However, any theoretical model of the home bias should also reproduce the fact that the relative importance of risk sharing channels does not vary over time. This may, for example, be an important restriction on transaction-cost based explanations of the home bias as the presence of (fixed) transactions costs may well imply that the relative roles of intertemporal smoothing and cross-sectional insurance vary with the forecast horizon.

Another issue that our results may raise is why permanent and transitory shocks are insured in such different ways. Apparently, permanent shocks are much harder to insure internationally than are transitory shocks. Why this should be the case is not immediately clear but it is what the data tell us. We plan to address this question in future research.

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**Data Appendix**

All data are from the Penn World Tables release 5.6

| Variable | PWT-code | Line# |
|---|---|---|
| GDP | cgdp | 9 |
| GNP | rgnp | 27 |
| C | cc | 10 |
| Population | pop | 1 |

The sample range is 1970-90.

List of Countries:

1. Canada, 2. 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. Norway, 17. Portugal, 18. Spain, 19. Sweden, 20. Switzerland, 21. United Kingdom, 22. Australia, 23. New Zealand

G7: countries #1,2,3,8,9,13,21

EU 15: countries #4,5,6,7,8,9,10,11,12,13,14,15,17,18,19,21

EU core: countries #4,5,8,9,14,15

## Technical Appendix

Anderson and Hsiao (1982) were the first authors to present a solution to the problem of estimating dynamic panel data models. For the case of a univariate AR(1), they suggested first differencing the model to get rid of the country-specific effects $f_i$ and then using $(\mathbf{\Delta X}_{i,t-2} - \mathbf{\Delta X}_{i,t-3})$ or simply $\mathbf{\Delta X}_{i,t-2}$ as an instrument for $(\mathbf{\Delta X}_{i,t-1} - \mathbf{\Delta X}_{i,t-2})$. These instruments will not be correlated with $(\mathbf{u}_{i,t} - \mathbf{u}_{i,t-1})$. This instrumental variables (IV) estimation method leads to consistent but not necessarily efficient estimates of the parameters of the model. In the sequel, several other studies (e.g. Arellano and Bond, 1991) suggested instruments leading to more efficient estimates. The above-mentioned problems are not specific to VARs.

In a landmark paper, Holtz-Eakin, Newey and Rosen (1988) - HNR for short - explained how to estimate VARs in a panel framework and proposed an IV type estimation procedure which we will now briefly explain.[5]

The specification of (8) as a projection implies that the error term $\mathbf{u}_{it}$ satisfies the orthogonality condition

$$E[\mathbf{\Delta X}'_{is}\mathbf{u}_{it}] = E[\mathbf{f}'_i\mathbf{u}_{it}] \qquad s < t \tag{13}$$

We can exploit these orthogonality conditions to identify the parameters of the model. Taking first differences on (8), we obtain

$$\begin{aligned} \mathbf{\Delta X}_{it} - \mathbf{\Delta X}_{it-1} &= \mathbf{\Lambda}_t + \sum_{l=1}^{p} \Phi_l(\mathbf{\Delta X}_{i,t-l} - \mathbf{\Delta X}_{i,t-l-1}) + \mathbf{v}_{it} & (14)\\ i &= 1, ..., K; t = p+2, ..., T & (15) \end{aligned}$$

where

$$\begin{aligned} \mathbf{\Lambda}_t &= \lambda_t - \lambda_{t-1} & (16)\\ \mathbf{v}_{it} &= \mathbf{u}_{it} - \mathbf{u}_{i,t-1} \end{aligned}$$

[5]Holtz-Eakin et al. deal with the more general case of an interacted country-specific and time effect $\lambda_t f_i$ and with time-varying coefficients.

We will now discuss identification of the parameters of the transformed equation (14) and then see how the original parameters can be recovered.

The orthogonality conditions of equation (13) imply that the error term of the transformed equation (14) satisfies the orthogonality condition

$$E[\mathbf{\Delta X}'_{is}\mathbf{v}_{it}] = E[\mathbf{f}'_i\mathbf{v}_{it}] \qquad s < t-1 \tag{17}$$

Therefore,

$$\mathbf{Z}_{it} = [\mathbf{e}, (\mathbf{\Delta X}_{i,t-2} - \mathbf{\Delta X}_{i,t-3}), (\mathbf{\Delta X}_{i,t-3} - \mathbf{\Delta X}_{i,t-4}), .., (\mathbf{\Delta X}_{i2} - \mathbf{\Delta X}_{i1})]$$

qualify as instrumental variables. The original parameters are identified if $T \geq p+3$.[6] Note that the number of instruments increases with $t$. Thus, the HNR estimator is more efficient than an IV estimator based on once-lagged endogenous variables alone (as in Anderson and Hsiao, 1982).

Estimation yields the coefficients $[\mathbf{\Phi}_1, ..., \mathbf{\Phi}_p]$, and we can calculate the variance-covariance matrix $\mathbf{\Omega}^*$ of the transformed system. Using (16), we are able to recover the variance-covariance matrix of the original system. The estimated coefficients $[\mathbf{\Phi}_1, ..., \mathbf{\Phi}_p]$ can be used to obtain the coefficient matrices $\mathbf{C}_l$ of the moving average representation. Finally, we can compute the mean squared prediction error using (6) from which the results in the main text follow immediately.

[6] Alternatively, following Arellano (1989) we used "level" values $Z_{it} = [\mathbf{e}, \mathbf{\Delta X}_{i,t-2}, \mathbf{\Delta X}_{i,t-3}, ..., \mathbf{\Delta X}_{i1}]$ as instruments in which case we also gain one more period for estimation because in this case identification only requires $T \geq p+2$. The results, however are very similar which we consider a robustness check of our empirical strategy.

**Table 1: Risk Sharing at various horizons**

| Country group | | | Forecast Horizon in years 1 | 3 | Unconditional Model |
|---|---|---|---|---|---|
| a) | All | $\beta_C$ | 0.09 (0.004) | 0.09 (0.004) | 0.001 (0.03) |
| | | $\beta_S$ | 0.29 (0.002) | 0.29 (0.002) | 0.32 (0.14) |
| b) | G7 | $\beta_C$ | −0.03 (0.01) | −0.03 (0.01) | 0.01 (0.03) |
| | | $\beta_S$ | 0.51 (0.04) | 0.51 (0.04) | 0.33 (0.13) |
| c) | EU15 | $\beta_C$ | 0.004 (0.01) | 0.004 (0.01) | −0.02 (0.04) |
| | | $\beta_S$ | 0.19 (0.05) | 0.19 (0.05) | 0.21 (0.18) |
| d) | EU core | $\beta_C$ | −0.02 (0.02) | −0.02 (0.02) | −0.02 (0.04) |
| | | $\beta_S$ | 0.33 (0.09) | 0.33 (0.09) | 0.34 (0.18) |

Values in parentheses are standard deviations.
For the unconditional model these are asymptotic whereas for the dynamic model they have been obtained using 100 bootstrap replications.

**Table 2: Share of first principal component and correlation with GDP shocks**

| | Variance explained by 1. PC | Correlation of 1. PC with $\varepsilon_{i1t}$ | $\varepsilon_{i2t}$ | $\varepsilon_{i3t}$ |
|---|---|---|---|---|
| All | 91% | 0.99 | -0.09 | 0.00 |
| G7 | 94% | 0.99 | -0.13 | 0.00 |
| Euro 15 | 77% | 0.99 | -0.13 | 0.00 |
| EU core | 78% | 0.97 | -0.23 | 0.00 |

**Table 3: Smoothing of permanent and transitory shocks**

| Country group | | | Type of shock | | Variance share |
|---|---|---|---|---|---|
| | | | permanent | transitory | of perm. shocks |
| a) | All | $\beta_C$ | 0.10 (0.005) | 0.03 (0.01) | |
| | | $\beta_S$ | 0.11 (0.03) | 0.91 (0.40) | 0.77 (0.03) |
| b) | G7 | $\beta_C$ | −0.05 (0.02) | −0.03 (0.03) | |
| | | $\beta_S$ | 0.36 (0.06) | 0.96 (0.32) | 0.72 (0.05) |
| c) | EU15 | $\beta_C$ | −0.007 (0.02) | 0.03 (0.03) | |
| | | $\beta_S$ | 0.54 (0.07) | −0.62 (0.38) | 0.70 (0.04) |
| d) | EU core | $\beta_C$ | −0.15 (0.05) | 0.11 (0.05) | |
| | | $\beta_S$ | 0.04 (0.19) | 0.62 (0.26) | 0.49 (0.09) |

Standard errors (in parentheses) were obtained from 100 bootstrap replications of the model

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