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Measuring International Risk-Sharing: Theoretical Issues and Empirical Evidence from OECD Countries

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Abstract

Whether financial market integration raised global insurance is a crucial, still open issue. All empirical methods to measure cross-border risk-sharing are based on the implicit assumption that international prices do not fluctuate in response to business cycle shocks. This paper shows that these methods can be completely misleading in the presence of large fluctuations in international prices as those observed in the data. I then propose a new empirical method that is immune from this issue. The risk-sharing inefficiency between two countries is measured by the wedge between their Stochastic Discount Factors (SDFs). This measure is a proxy for the welfare losses created by imperfect insurance. Welfare losses can be attributed either to the strength of uninsurable shocks (the extent of risk to be pooled) or to the degree of insurance against different sources of risk. The method is applied to study the evolution of risk-sharing between the US and OECD countries, assuming either constant or time-varying risk-aversion. The degree of insurance is found to have improved over time only for some countries and only if SDFs are estimated assuming time-varying risk-aversion. The results are also informative on the implications of different macro models for international risk. When confronted with the data, standard open-macro models (featuring constant risk-aversion) imply that nominal exchange rate fluctuations do not contain wealth divergences across countries, but rather represent an important source of risk. Time-varying risk-aversion instead implies that limiting welfare losses from imperfect risk-sharing requires reducing the volatility of macro fundamentals.

1 Introduction

The last decades have been characterized by a fast process of international financial integration. There is quite a generalized agreement that removing barriers to international asset trade increased the volume of financial resources flowing across borders. Yet whether financial integration effectively raised global insurance and welfare, is still an open issue. The literature soundly rejected the hypothesis of full insurance against national income shocks. However, results about the actual degree of international risk-sharing and its evolution over time are still not conclusive. One reason is that several empirical methods

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are well-designed to test the null hypothesis of full risk-sharing, but once this null is rejected they cannot (by construction) quantify the actual lack of insurance.[1] On the other hand, the methods that were explicitly designed to measure the degree of insurance, are all based on the implicit assumption that international prices do not fluctuate in response to business cycle shocks.[2] Viani (2010) studied from a theoretical point of view the problems that can be created by this assumption. Namely, postulating no price fluctuations implies that international financial markets are efficient if they channel resources to countries that (absent asset trade) would be made relatively poorer by national shocks (net of physical capital accumulation). Against this standard paradigm, Viani (2010) shows that the interaction between relative price fluctuations and cross-border financial flows is crucial to derive testable conditions for cross-border risk-sharing. In particular, if receiving an inflow from abroad lowers the international price of a country's output, efficient financial markets should channel resources "upstream", from poorer to richer countries. These financial flows are optimal, yet they contradict the condition of efficiency on which standard empirical methods are based.

From an empirical perspective, if the assumption of no relative price fluctuations – at the basis of virtually all empirical tests – might be a relatively innocuous approximation when measuring insurance among different regions within a country (a tasks for which some of these methods were originally designed), it could be instead problematic when quantifying cross-border risk-sharing. Indeed, deviations from Purchasing Power Parity across countries have been estimated to be large and persistent.

In this paper I assess the performance of standard methods in a simulated two-country DSGE model with endogenous portfolio diversification, in which relative price fluctuations can transfer purchasing power across countries. Results show that standard methods may be completely misleading in the presence of large movements in international prices as those observed in the data, as the benchmark of efficiency on which they are based is invalidated.

I then propose a new empirical method to quantify cross-border risk-sharing, which is immune from these issues. The method consists in estimating the risk-sharing inefficiency between two countries (contingent on the shocks that hit them during a certain period) using the wedge between their Stochastic Discount Factors (SDFs). This measure represents a proxy for the aggregate welfare losses created by imperfect insurance. Welfare losses can then be attributed either to the strength of uninsurable shocks (the extent of risk to be shared) or to the degree of insurance against different sources of risk.

The method is applied to study the evolution of risk-sharing between the US and industrialized OECD economies over the last forty years, assuming either constant or time-varying risk-aversion. I find that the degree of insurance has improved over time only for some countries and only if SDFs are estimated assuming time-varying risk-aversion. The results are also informative on the implications of different macro models for international risk. When confronted with the data, standard open-macro models (postulating constant risk-aversion) imply that nominal exchange rate fluctuations do not contain wealth divergences across countries, but rather represent an important source of risk. Time-varying risk-aversion instead implies that limiting welfare losses from imperfect risk-sharing requires reducing the volatility of macro fundamentals.

The first part of the paper explores the issues to which standard empirical methods – based on the assumption of no international price fluctuations – may be subject. To this aim, I write a simple two-country two-good DSGE model in which relative price

[1] See for instance Kollmann (1995) and Ravn (2001).

[2] Among these, the methods developed by Asdrubali, Sorensen and Yosha (1996), Obstfeld (1993), Brandt, Cochrane and Santa Clara (2006), Flood, Marion and Matsumoto (2008).

movements can transfer purchasing power across countries, and simulate it under three different asset structures – Financial autarky, trade in uncontingent bonds, and trade in two equities. By construction Financial autarky gives less or equal insurance against national supply shocks than the other asset structures. Trade in uncontingent bonds ranks second, while trade in two equities gives perfect risk-sharing against national disturbances. I apply different empirical methods (respectively, from Asdrubali, Sorensen and Yosha (1996), Obstfeld (1993), Brandt, Cochrane and Santa Clara (2006), and Flood, Marion and Matsumoto (2008)) to the simulated model in order to check if they identify correctly full risk-sharing in the two-equity case, and if they capture the correct ranking of insurance among asset structures. The results unveil three typical issues that may affect standard methodologies. First, the relative value of GDPs commonly used as the independent variable in risk-sharing regressions, does not necessarily identify which country is made relatively richer/poorer by national shocks, which makes existing methods potentially misleading. Second, the interaction between relative prices and financial flows may require efficient financial markets to channel resources "upstream", to relatively richer countries – a finding that is typically interpreted as low insurance by existing methods. Finally, the efficient consumption allocation can be characterized only in conjuncture with relative price movements: with full insurance national consumption can fall after a negative income shock, provided consumers are compensated by an appreciation of their currency – a possibility that is not considered by standard benchmarks of efficiency. These issues can have potentially severe consequences. Not only they can lead to reject full insurance also when the "true" model features effectively complete asset markets, but they can also lead to estimate a higher degree of insurance under, say, Financial autarky than in the case of complete markets. In a world in which relative price fluctuations are realistically large and sustained, the benchmarks of efficiency derived in one-good frameworks and on which standard methods are based, do not seem to be valid anymore. Most existing empirical methods may then be easily misleading.

In the remaining of the paper I propose a new method to quantify cross-border risk-sharing. Macro theory suggests that perfect risk-sharing equalizes the Stochastic Discount Factors (SDFs) of any two economies in any state of the world. Therefore the risk-sharing inefficiency between countries (or country aggregates) can be measured by the wedge between their SDFs. Using a simple theoretical framework I prove that this measure of cross-border risk-sharing, the *gap*, reflects changes in countries' relative wealth created by uninsurable shocks. Under certain conditions it also maps into aggregate welfare losses due to imperfect insurance. I show how its variance can be used to quantify the *contingent risk-sharing inefficiency* between two countries over a certain time horizon. By regressing the time-series of the *gap* estimated from the data on different sources of macroeconomic risk (national GDP volatility, government spending, nominal exchange rate fluctuations), we can test whether households in different countries are fully insured against these risks. Falls in the contingent inefficiency over time can be attributed either to reductions in the strength of uninsurable shocks (the extent of risk to be shared) or to improvements in the degree of insurance against different sources of risk. The performance of this method is checked using the simulated DSGE model. Results show that the present methodology properly identifies full-risk sharing and captures the correct ranking of insurance across different asset trade regimes.

Finally, the method is applied to study the evolution of risk-sharing between the US and 16 industrialized OECD countries over the period 1970-2008. Stochastic Discount Factors are estimated using two alternative strategies. First, I rely on the assumptions of standard open-macro models, and assume constant relative risk-aversion. The degree of risk-aversion is then estimated from the data using non-linear GMM.[3] The second strategy consists in following Campbell and Cochrane (1999) and postulating a utility function with time-varying risk-aversion. I adopt the parametrization that was used in that paper to replicate the equity-premium puzzle and to capture the history of US stock prices.

I find that the degree of insurance against national idiosyncratic shocks has improved over time only for some countries and only if SDFs are estimated assuming time-varying risk-aversion. Different macro models are shown to have different implications for international risk. When SDFs are estimated relying on the assumptions of the standard macro model (constant risk-aversion), I find that nominal exchange rate volatility represented an important source of macroeconomic risk, whose decrease led to a significant reduction in welfare losses from imperfect insurance. Thus when confronted with the data, standard open-macro models imply that nominal exchange rate fluctuations do not contain wealth divergences across countries. Instead, a reduction in nominal exchange rate volatility can reduce substantially the risk to be shared internationally and increase global welfare. This result bears interesting implications for the literature on optimal exchange rate regimes. While in a well-know analysis Baxter and Stockman (1988) found that the only change in the behaviour of macro aggregates caused by the adoption of a fixed exchange rates regime, is a reduction in exchange rate volatility, our results suggest that a fall in this volatility can indeed reduce substantially cross-border risk-sharing inefficiencies.

When SDFs are estimated assuming time-varying risk-aversion, I find that the fall in the volatility of macro aggregates – mainly GDP of both the US and OECD countries, and its components – have reduced significantly contingent risk-sharing inefficiencies. The Great Moderation, the lower GDP volatility experienced in the last decades by both the US and most industrialized countries, seems to have increased global welfare by reducing inefficiencies due to imperfect risk-sharing. Thus time-varying risk-aversion implies that limiting welfare losses from imperfect risk-sharing requires reducing the volatility of macro fundamentals.

This paper is related to Viani (2010) which questions the validity of the theoretical efficiency condition for financial markets at the basis of most empirical methods. The present paper is the natural follow up of that analysis, as it investigates the implications of those findings for empirical methods. This paper is also related to the vast literature on empirical risk-sharing measurement. Asdrubali, Sorensen and Yosha (1996), Obstfeld (1993), Brandt, Cochrane and Santa Clara (2006), and Flood, Marion and Matsumoto (2008) are only a non-exhaustive list. Virtually all these methods are implicitly based on the assumption of no relative price fluctuations.

To my knowledge this paper is the first one to assess the performance of existing empirical methods using a simulated model, and to investigate the problems that can be caused by the presence of relative price fluctuations. It is also the first one to propose a theoretically-consistent measure of cross-border risk-sharing inefficiencies that gives an indication of the welfare losses created by imperfect insurance and does not rely on restrictive assumptions about relative price movements.

The rest of the paper is organized as follows. The next section studies the performance of standard empirical methods through simulation analysis. Section 3 describes the new method and its theoretical foundations, and assesses its performance in simulation exercises. Section 4 discusses the empirical strategy. Section 5 shows the results from applying this method to the US and OECD countries. Section 6 draws some conclusions.

[3] GMM estimation also allows verifying that the functional form assumed is not rejected by the data.

2 Assessing the performance of standard methods

The purpose of this section is to investigate which problems can be caused to existing empirical methods by the presence of relative price fluctuations. I write a simple model in which both relative price fluctuations and financial flows can transfer purchasing power across countries, and simulate it under three different asset structures – Financial autarky, trade in uncontingent bonds, and trade in two equities. The risk-sharing properties of these asset trade regimes have been studied in Viani (2010). Financial autarky gives less or equal insurance against national supply shocks than the other asset structures. Trade in uncontingent bonds ranks second, while trade in two equities gives perfect risk-sharing against national disturbances. I apply different empirical methods to the simulated model to check if they identify correctly full risk-sharing in the two-equities case, and if they capture the correct ranking of insurance among asset structures. I report here the results relative to the method of Asdrubali, Sorensen and Yosha (1996). I use them to illustrate the problems that can be caused to existing empirical frameworks by the presence of relative price fluctuations and by their interaction with cross-border financial flows. Results from applying other methods (Obstfeld (1993), Brandt, Cochrane and Santa Clara (2006), Flood, Marion and Matsumoto (2008)) are reported in Appendix A.

2.1 The model

The model is the two-country two-good DSGE framework with endowment shocks and home bias in consumption studied in Viani (2010).[4] Its core structure is the simplest one that generates a general equilibrium interaction between the two channels of international insurance, relative price fluctuations and cross-border financial flows.

The model consists of two countries, Home and Foreign, each inhabited by a representative household. Countries are specialized in the production of different goods. Households in country H receive utility from consuming a bundle made up of the foreign and the domestic good, according to a CES aggregator

$$C_t = \left[(\delta)^{1/\theta} (C_{ht})^{\frac{\theta-1}{\theta}} + (1-\delta)^{1/\theta} (C_{ft})^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{\theta-1}} ; \theta > 0, \delta > 1/2 \tag{1}$$

where C_{jt} denotes consumption of the good produced in country j and δ is a parameter capturing home bias in consumption. θ is the elasticity of substitution between H and F-produced goods; in this model it coincides with the trade elasticity. The F consumption bundle is analogously defined. Starred variables denote the corresponding quantities consumed by Foreign agents. Goods are freely tradable but not storable. The period utility of both agents depends on current consumption only and is a Constant Relative Risk Aversion function with risk aversion coefficient ρ.

In each period H and F households receive a stochastic endowment according to the process

$$\log(Y_{jt}) = \zeta \log(Y_{jt-1}) + \varepsilon_{jt}$$

where Y_{jt} denotes the endowment received by consumer j and $\varepsilon_{jt} \sim iid(0, \sigma^{\varepsilon})$.

I assume that the law of one price holds and that the nominal exchange rate is constant and equal to one for simplicity. Due to home bias in consumption Purchasing Power Parity does not hold outside a symmetric Steady State, and the two price indexes P_t and $P_t^{\star}$ are tied by the following condition defining the real exchange rate Q_t

[4] The model is closely related to frameworks employed in Cole and Obstfeld (1991), Kollmann (2006), and Corsetti et al (2008).

$$Q_t \equiv \frac{P_t^\star}{P_t}$$

Terms of trade are defined as the ratio of the price of H imports and exports

$$tot_t \equiv \frac{P_{ft}}{P_{ht}}$$

where P_{ht}, and P_{ft} denote respectively the price of H and F-produced goods.

Standard cost minimization delivers H households' demand for the H and the F variety

$$C_{ht} = \delta \left(\frac{P_{ht}}{P_t} \right)^{-\theta} C_t$$

$$C_{ft} = (1-\delta) \left(\frac{P_{ft}}{P_t} \right)^{-\theta} C_t$$

Analogous conditions hold for F agents. Good market clearing conditions read

$$Y_{ht} = C_{ht} + C_{ht}^\star$$

$$Y_{ft} = C_{ft} + C_{ft}^\star$$

The model is loglinearized around a symmetric Steady State in which countries' wealth is assumed to be equalized, endogenous variables are constant and exogenous ones are equal to their mean values.

I simulate the model under three different asset trade regimes. Households' budget constraints are pinned down by the specific asset structure assumed.[5]

2.2 Asset structure 1: Financial autarky

Under Financial autarky relative price fluctuations represent the only channel of cross-border insurance. Households have no means of smoothing consumption over time and in each period their consumption must equal the value of their income. H and F budget constraints read

$$C_t P_t = P_{ht} Y_{ht}$$

$$C_t^\star P_t^\star = P_{ft} Y_{ft}$$

Viani (2010) shows that in this model Financial autarky gives less or equal insurance against national shocks than the other two asset trade regimes.

[5] In order to focus squarely on the problems caused to empirical methods by the presence of relative price fluctuations, we will abstract from the following issues, well-known to the insurance literature: taste shocks, measurement errors, lack of capital gains/losses in National Accounts data, econometric issues due to panel estimation.

2.3 Asset structure 2: Trade in uncontingent bonds

If agents can trade in uncontingent bonds paying 1 unit of the H consumption bundle in every state-of-the-world, the budget constraint of H consumers reads

$$P_t^B B_{t+1} = \frac{P_{ht} Y_{ht}}{P_t} - C_t + B_t$$

where B_{t+1} denotes bonds purchased in period t, P_t^B their unitary price in terms of the H consumption bundle. Bonds are assumed to be in zero net supply. H households' inter-temporal problem is given by

$$\max_{\{C_t, B_{t+1}, \nu_t\}} . : L = E_0 \left\{ \sum_{t=0}^{\infty} \beta^t \left[\frac{(C_t)^{1-\rho}}{1-\rho} + \nu_t \left(P_{ht} Y_{ht} + P_t B_t - C_t P_t - P_t^B P_t B_{t+1} \right) \right] \right\}$$

H agent's first order conditions read

$$(C_t)^{-\rho} = \nu_t P_t$$

$$P_t^B = \beta E_t \left\{ \frac{\nu_{t+1}}{\nu_t} \frac{P_{t+1}}{P_t} \right\}$$

In the symmetric Steady State $B = B^\star = 0$.

Viani (2010) shows that trade in uncontingent bonds gives the second-best degree of insurance among the three asset structures considered.

2.4 Asset structure 3: Trade in two equities (effectively complete markets)

Assume households can trade in two equities, H and F, representing respectively a claim to the H and F country's endowment. The total quantity of each equity is normalized to 1. S_h, S_f, S_h^*, and S_f^* denote the fraction of H and F equities owned respectively by H and F consumers. Due to the normalization, the owner of S_h equities receives a share S_h of the H endowment. Equities real returns, expressed in terms of the H consumption good, are given by

$$R_t = \frac{(P_{ht}/P_t) Y_{ht} + Z_t}{Z_{t-1}}$$

$$R_t^* = \frac{(P_{ft}/P_t) Y_{ft} + Z_t^*}{Z_{t-1}^*}$$

where Z and Z^*denote the real price (in terms of the H consumption bundle) of the two assets.

The budget constraint of H households reads

$$NFA_t = NFA_{t-1} + \frac{Y_{ht} P_{ht}}{P_t} - C_t + R_t \cdot (S_{ht} - 1) + R_t^* S_{ft}$$

where $NFA_t \equiv [(S_{h,t+1} - 1) \cdot Z_t + S_{f,t+1} \cdot Z_t^*]$ denotes H net foreign assets.

Euler equations and asset market clearing conditions are given by

$$(C_t)^{-\rho} = \beta E_t \left\{ (C_{t+1})^{-\rho} \cdot R_{t+1} \right\} = \beta E_t \left\{ (C_{t+1})^{-\rho} \cdot R_{t+1}^* \right\}$$

$$\frac{\left(C_t^{\star}\right)^{-\rho}}{Q_t}=\beta E_t\left\{\frac{\left(C_{t+1}^{\star}\right)^{-\rho}}{Q_{t+1}}\cdot R_{t+1}\right\}=\beta E_t\left\{\frac{\left(C_{t+1}^{\star}\right)^{-\rho}}{Q_{t+1}}\cdot R_{t+1}^{*}\right\}$$

In the Steady State $NFA=NFA^{\star}=0$.

As shown in Kollmann (2006) and Viani (2010), trade in two equities makes asset markets effectively complete and provides perfect risk-sharing.

2.5 Parametrization

I simulate the behaviour of the model for 1000 periods assuming the following parameter values

Table 1: Model parametrization

H consumption home bias	δ	0,9
F consumption home bias	$\delta^{\star}$	0,9
H discount factor	β	1/1.01
F discount factor	$\beta^{\star}$	1/1.01
Risk-aversion coefficient	ρ	2
H shocks persistence	ζ	0,95
F shocks persistence	$\zeta^{\star}$	0,95

I also assume $\varepsilon_t \sim N(0,1)$, $\varepsilon_t^{\star} \sim N(0,1)$, and $Cov(\varepsilon_t, \varepsilon_t)=0$.

2.6 The performance of empirical methods: Asdrubali, Sorensen and Yosha (1996)

2.6.1 The methodology

The method of Asdrubali, Sorensen and Yosha (1996) (ASY henceforth) was originally developed to estimate risk-sharing among US states. Yet most subsequent applications used it to estimate cross-border insurance. The methodology assumes no relative prices and relies on the following identity[6]

$$C=GDP+NFI-NSAV \tag{2}$$

where NFI and $NSAV$ denote respectively net financial income and net savings. From (2), full insurance – C and Y being uncorrelated, according to ASY – requires a country experiencing a negative output shock (getting poorer absent asset trade) to receive income

[6]Equation (2) implies that consumption and output have the same composition, hence the same price.

from abroad (either from portfolio returns or from external borrowing – net of physical capital accumulation). The method consists in estimating

$$\Delta \log C_t - \Delta \log C_t^\star = \alpha + \beta \cdot (\log Y_t - \log Y_t^\star) + \varepsilon_t \tag{3}$$

where Δ denotes the first difference of a variable and β is interpreted as a measure of the uninsured effects on consumption of GDP fluctuations.

Moreover ASY show that the decomposition in (2) allows to quantify the fraction of output shocks that is absorbed through different channels of insurance by panel estimating

$$-\Delta \log NFI_t \equiv (\Delta \log GDP_t - \Delta \log GNP_t) = \alpha_{f,t} + \beta_f \cdot \Delta \log GDP_t + \varepsilon_{ft} \tag{4}$$

$$\Delta \log NSAV_t \equiv (\Delta \log GNP_t - \Delta \log C_t) = \alpha_{s,t} + \beta_s \cdot \Delta \log GDP_t + \varepsilon_{st}$$

$$\Delta \log C_t = \alpha_{u,t} + \beta_u \cdot \Delta \log GDP_t + \varepsilon_{ut}$$

where β_f is interpreted as the fraction of output shock that is absorbed through factor income flows, β_s is the share absorbed through consumption smoothing, and β_u the fraction left uninsured. Time fixed effects in (4) are meant to capture the effects of aggregate shocks. In an OLS framework, we can replace time fixed effects by adding Foreign GDP as a regressor, and estimate

$$(\Delta \log GDP_t - \Delta \log GNP_t) = \alpha_f + \beta_f \cdot \Delta \log GDP_t + \gamma_f \cdot \log GDP_t^\star + \varepsilon_{ft} \tag{5}$$

$$(\Delta \log GNP_t - \Delta \log C_t) = \alpha_s + \beta_s \cdot \Delta \log GDP_t + \gamma_s \cdot \log GDP_t^\star + \varepsilon_{st}$$

$$\Delta \log C_t = \alpha_u + \beta_u \cdot \Delta \log GDP_t + \gamma_u \cdot \log GDP_t^\star + \varepsilon_{ut}$$

Notice that under the maintained assumption of no relative prices, equations (5) imply the same view of optimal financial flows underlying (2): efficient financial markets should channel resources to the countries that (absent asset trade) would have been made relatively poorer by national shocks. Slopes in the interval $[0, 1)$ are interpreted as a relatively high degree of insurance. Instead, $\beta_u > 1$, $\beta_f < 0$, and $\beta_s < 0$ indicate that financial markets are highly inefficient. Kalemli-Ozcan et al. (2004) estimate $\beta_f < 0$ among European countries during the Nineties, and interpret this finding as signaling bad insurance from cross-border ownership of securities.

A first adjustment to account for price fluctuations Since this method had been originally developed to estimate risk-sharing among different regions within a country and abstracted completely from relative price movements, most studies estimating (3) and (5) in an international setup implicitly modified identity (2) to account for international price fluctuations. If relative prices are well-defined, (2) becomes

$$C = \frac{P_y \cdot GDP}{P} + NFI - NSAV \tag{6}$$

where P and P_y are respectively the prices of consumption and output, and net financial income and net savings are expressed in terms of real consumption. Thus most studies deflate national output with the local CPI, so as to express all variables in terms of the country's consumption.[7] In all the estimations below I will assume this correction, and deflate output using national CPIs. Appendix B discusses the consequences of not implementing this change and shows the corresponding results.

Notice that by subtracting the national identities (6) relative to countries H and F we can write their consumption differential as[8]

$$
\begin{aligned}
(\Delta \log C_t - \Delta \log C^\star) &= \underbrace{(\Delta \log GDP_t - \Delta \log GDP_t^\star) + \omega \Delta \log TOT_t}_{\Delta \log v_t} \\
&\quad - \phi \Delta \log NSAV_t + \kappa \Delta \log NFI_t
\end{aligned}
$$

where TOT_t are country H terms of trade, ω, ϕ, and κ are positive constants, and ν_t is the real relative value of the H and the F endowments. Thus, running (3) with properly deflated output is equivalent to estimating

$$
(\Delta \log C_t - \Delta \log C^\star) = \alpha + \beta \cdot \Delta \log v_t + \varepsilon_t
$$

The impact on consumption differentials of relative price fluctuations is included in the independent variable, ν_t. Therefore, according to the ASY method, we can interpret β as the fraction of shocks that is *not* insured through financial markets.

2.6.2 Estimation results: The problems caused by relative price fluctuations

Table 2: ASY method
OLS estimation of the fraction of shocks left uninsured

β (fraction uninsured)	$\theta = 0.3$	$\theta = 0.6$	$\theta = 1.2$
Financial autarky	1	1	1
Bond	1,74	1,190	0,73
Complete markets	1,75	1,191	0,72

Table 2 reports the estimation results for specification (3). Standard errors are negligible and are not reported. The estimation is carried out for three different values of the trade elasticity θ. It can be shown that in the present model these values correspond to different kinds of interactions between price fluctuations and financial flows. For $\theta = 0.3$ receiving a financial inflow from abroad lowers the relative price of a country's output (prices and flows are substitutes in providing insurance). For $\theta > 0.3$ the same transfer raises the international price of a country's output (the two insurance channels, flows and prices, are complements).[9]

[7] See for example Sorensen and Yosha (1998), Del Negro (2000), and Kalemli-Ozcan et al. (2004). Sorensen and Yosha (1998) and Kalemli-Ozcan et al. (2004) motivate the need for this adjustment.

[8] All aggregates are expressed in terms of Home consumption.

[9] See Viani (2010).

The estimations identify correctly Financial autarky as the case in which financial markets do not absorb any fraction of national disturbances ($\beta = 1$). Notice however that in this model the absence of asset trade has a very different impact on consumers' welfare for different values of the trade elasticity. Viani (2010) shows that under Financial autarky cross-border insurance from relative price fluctuations is very high for $\theta = 1.2$, while terms of trade movements enlarge the risk-sharing inefficiency that would have aroused at constant prices for $\theta = 0.3$. The β coefficients estimated by the ASY method for the Financial autarky specification do not give any indication of the welfare losses due to the absence of financial markets.

For $\theta = 1.2$ the estimated coefficients identify the correct ranking of insurance across asset trade regimes ($\beta^{CM} < \beta^{B} < \beta^{FA}$). However, full insurance is rejected in the specification with complete markets – more than 70% of the shocks is estimated to be left uninsured. For $\theta = 0.6$ and $\theta = 0.3$ estimating (3) leads to reject the null of full insurance in the complete markets case. Moreover the coefficients do not even identify the correct ranking of insurance across asset structures. The coefficients estimated for the bond and the complete markets case exceed unity, signaling worse insurance than under Financial autarky.

Notice that from an econometric point of view all the coefficients are estimated correctly. The failure to identify full risk-sharing and the correct ranking of insurance reflects the following issues, which concern the *interpretation* of regression (3)

Issue 1: The consumption allocation corresponding to full risk-sharing depends on relative prices According to the ASY method, an allocation delivers full insurance if consumption differentials and GDP differentials are uncorrelated. However from a theoretical point of view, when relative prices are well-defined, consumption and GDP differential may well be correlated in the full risk-sharing allocation. Macroeconomic theory suggests that with full insurance agents' Stochastic Discount Factors should be equalized at any time. Viani (2010) shows that in the present model this condition implies that the allocation achieved in an equilibrium with complete asset markets, should be characterized *both* by relative consumption and relative prices, according to

$$\frac{U_{C,t}}{U_{C,t-1}} \frac{RER_t}{RER_{t-1}} = \frac{U^{\star}_{C,t}}{U^{\star}_{C,t-1}}, \forall t$$

where U_C and $U^{\star}_C$ denote respectively the marginal utility of consumption in the H and the F country, and RER is the real exchange rate between the two currencies. In terms of growth rates, and assuming CRRA utility function with risk-aversion parameter ρ

$$\Delta \log RER = \rho \cdot (\Delta \log C_t - \Delta \log C^{\star}_t)$$

In the full insurance allocation consumption in the Home country can fall in response to a negative income shock, provided H consumers are compensated by an appreciation of their currency. The benchmark of risk-sharing underlying equation (3) does not take into account the role played by relative prices in characterizing the full insurance allocation. This is why the ASY method estimates $\beta \neq 0$ even in the model with complete asset markets.

Issue 2: The relative value of endowments does not reflect relative wealth effects The ASY method regresses consumption differentials on the relative value of endowments. The presumption is that an increase in the relative value of the H endowment (the Home country getting relatively richer in equilibrium) should be followed by a financial transfer to the Foreign country if financial markets foster insurance. However, the relative value of endowments does not necessarily indicate which country is made relatively poorer by national shocks. It is possible to show that for $\theta = 0.6$ an increase in the Home endowment raises its relative value under any asset trade regime. But due to a strong terms of trade depreciation, the shock would make Home consumers relatively poorer absent asset trade.[10] Thus, when trade in assets is introduced, a rise in the Home endowment is optimally followed by a transfer of financial resources to this country, both in the model with trade in two equities and with trade in bonds. For $\theta = 0.6$ estimating $\beta > 1$ (and a larger β in the case of complete markets) reflects resources flowing to the Home country after an increase in the value of its GDP. Contrary to the interpretation of ASY, these flows are optimal and signal a high degree of insurance. What the method does not consider is that the relative value of endowments does not reflect relative wealth effects.

Issue 3: Efficient financial markets may channel resources "upstream" It can be proved that when flows and prices are substitutes (when they channel purchasing power in opposite directions) risk-sharing requires financial markets to channel resources to countries that (absent asset trade) would be made relatively *richer* by national shocks. This happens in our model for $\theta = 0.3$.[11] An increase in the relative value of the H endowment would make the Home country richer absent asset trade. Since flows and prices are substitutes, after the shock efficient financial flows channel resources "upstream", to the Home country, both in the model with trade in two equities and with trade in bonds. Estimating $\beta > 1$ in (3) indicates precisely "upstream" flows. Contrary to the standard interpretation of ASY, however, $\beta > 1$ corresponds to a high degree of insurance – perfect in the case of trade in two equities.

Table 3 reports the results from estimating specification (5). In the case of trade in two equities, full insurance is achieved through net factor income flows. The estimations should then find $\beta_f = 1$ for all the values of the trade elasticity. For $\theta = 1.2$ the estimation rejects full insurance through net factor income flows, but signals that a (small) share of supply shocks is absorbed through this insurance channel ($\beta > 0$). For $\theta = 0.6$ and $\theta = 0.3$, the estimation signals *negative* insurance from cross-border ownership of securities. We find similar results when estimating the share of shocks absorbed through consumption smoothing in the model with trade in uncontingent bonds.[12] The issues behind these results are the same discussed above.

2.6.3 Summarizing

Checking the performance of the ASY method in a model with relative price fluctuations has unveiled three issues that may affect this methodology. First, the relative value of endowments used as a regressor does not necessarily identify which country is made relatively richer/poorer by national shocks. Second, the interaction between relative prices and financial flows may require efficient financial markets to channel resources "upstream", to relatively richer countries. Finally, the efficient consumption allocation can be characterized only in conjuncture with relative price movements: with full insurance national consumption can fall after a negative income shock, provided consumers are compensated by an appreciation of their currency – a condition that is not considered by the ASY

[10] See Viani (2010) for a formal proof. $\theta = 0.6$ corresponds to $\tilde{\tau} < \theta < \tilde{\eta}$ in the notation of that paper. In this region of the trade elasticity domain, under Financial autarky an increase in the H endowment raises the shadow value of Home current income above the shadow value of Foreign current income – the shock makes the H consumer relatively poorer.

[11] See Viani (2010) for a formal proof. $\theta = 0.3$ corresponds to $\theta < \tilde{\tau}$ in that paper.

[12] The estimation signals (correctly) that no fraction of the shocks is absorbed through smoothing in the complete markets case. Clearly, estimating the share of shocks that is left uninsured yields the same results as for specification (3).

Table 3: ASY method
OLS estimation of shares absorbed through different channels

β**f (share absorbed via NFI)**	$\theta = 0.3$	$\theta = 0.6$	$\theta = 1.2$
Financial autarky	/	/	/
Bond	/	/	/
Complete markets	-0,38	-0,09	0,14
β**s (share absorbed via NSAV)**			
Financial autarky	/	/	/
Bond	-0,37	-0,09	0,13
Complete markets	0	0	0
β**u (share uninsured)**			
Financial autarky	1	1	1
Bond	1,37	1,095	0,863
Complete markets	1,38	1,0952	0,862

benchmark of efficiency. These problems can have potentially severe consequences. Not only they lead to reject full insurance also when the "true" model features effectively complete asset markets, but they can also lead to estimate a higher degree of insurance under, say, Financial autarky than in the case of complete markets.

In a world in which relative price fluctuations are large and sustained and contribute to transfer purchasing power across countries, the benchmarks of efficiency derived in one-good frameworks and on which standard methods are based, do not seem to be valid anymore. Looking at the correlation between relative consumption and relative output, that is focusing on the pattern of financial flows across countries as required by the ASY approach and most existing empirical methods, may then be easily misleading.

3 An insurance benchmark from macro theory

In this section I derive a measure of cross-border insurance valid also in presence of relative price fluctuations. This measure provides the theoretical foundations for a new empirical method to estimate cross-border risk-sharing. I apply the new method to the simulated model of section 2 in order to compare it with previous methodologies.

3.1 A measure of cross-border insurance

Assume the world consists of several countries, each inhabited by a representative agent. Asset markets are complete as agents can trade internationally in a full set of Arrow-Debreu securities. Focus on two countries, Home and Foreign. In what follows $W(s\prime|s)$ denotes the price in state-of-the-world s of an Arrow-Debreu security paying 1 unit of some numéraire good in the following period if state-of-the-world $s\prime$ realizes. $U_C(s)$ and $U_C(s\prime|s)$ are the marginal utilities of consumption of the H consumer, respectively, in state s and in state $s\prime$. $P(s)$ and $P(s\prime|s)$ are the prices of the H consumption good in state s and $s\prime$. β denotes the discount factor of the H consumer, $\pi(s\prime|s)$ the conditional probability of state $s\prime$ given s. The Euler Equation of the H consumer, regulating her purchase in state s of Arrow-Debreu securities paying 1 unit of the numéraire in state $s\prime$, reads

$$U_C(s)\frac{W(s\prime|s)}{P(s)} = \beta \cdot \pi(s\prime|s) \cdot U_C(s\prime|s)\frac{1}{P(s\prime|s)}, \forall s.s\prime \tag{7}$$

The H consumer buys securities until the marginal cost of purchasing one more asset (left hand side in equation (7)) equals its expected marginal benefit (right hand side).

The analogous condition for the F agent is given by

$$U_C^\star(s)\frac{W(s\prime|s)}{P^\star(s)\cdot\omega(s)} = \beta^\star \cdot \pi(s\prime|s) \cdot U_C^\star(s\prime|s)\frac{1}{P^\star(s\prime|s)\cdot\omega(s\prime|s)}, \forall s, s\prime \tag{8}$$

where starred variables denote Foreign aggregates, and ω is the nominal exchange rate between H and F currencies.

Combining (7) and (8) gives

$$\underbrace{\beta\frac{U_C(s\prime|s)}{U_C(s)}\frac{P(s)}{P(s\prime|s)}}_{m(s\prime|s)} = \underbrace{\beta^\star\frac{U_C^\star(s\prime|s)}{U_C^\star(s)}\frac{P^\star(s)}{P^\star(s\prime|s)}\frac{\varepsilon(s)}{\varepsilon(s\prime|s)}}_{m^\star(s\prime|s)} = \frac{W(s\prime|s)}{\pi(s\prime|s)}, \forall s, s\prime$$

where $m(s\prime|s)$ and $m^\star(s\prime|s)$ are the Stochastic Discount Factors (SDFs) of the H and the F consumer. Perfect cross-border risk-sharing against national income shocks (given by construction by complete asset markets) equates the SDFs of the two agents, for every states-of-the-world s and $s\prime$. Equivalently, in terms of time-dependent notation, perfect risk-sharing implies

$$m_t = m_t^\star, \forall t \tag{9}$$

If countries cannot pool risk efficiently the SDFs m and $m^\star$ need not be equalized: uninsurable shocks will drive a wedge between them in some states of the world. Therefore

$$m_t = \eta_t \cdot m_t^\star \tag{10}$$

where η_t represents the wedge between H and F SDFs created by uninsurable disturbances. Taking logarithms on both sides of (10) I define the measure of cross-border insurance, the *gap*, as

$$gap_t \equiv (\log \eta_t - \log 1) = \log m_t - \log m_t^\star \tag{11}$$

gap_t measures the percentual deviation of the wedge between H and F SDFs observed at time t from the wedge that would be observed if income risk was perfectly shared (i.e. if η was equal to 1). Therefore, by construction gap_t measures the percentual deviation from full insurance against income fluctuations observed at time t.

Appendix C shows that assuming frictionless trade in uncontingent bonds equalizes *expected* SDFs. In other words, self-insurance drives the conditional mean of the *gap* to zero

$$E_t\left(m_{t+1}\right) = E_t\left(m_{t+1}^\star\right) \Rightarrow E_t\left(gap_{t+1}\right) = 0 \tag{12}$$

It is well-known that if income risk is perfectly shared countries' relative wealth should stay constant over time. Using a simple two-country two-good DSGE model and following closely Viani (2010), Appendix C proves that the *gap* measures changes in countries' relative wealth created by uninsurable shocks. If consumers in different countries cannot fully insure against national shocks, idiosyncratic disturbances change their relative wealth. The sign of the *gap* arising in response to shocks indicates the new ranking of wealth, that is which country has been made relatively richer (poorer) by the shock. Its magnitude quantifies the change in countries' relative wealth.

Moreover, there exists a monotone relationship between the *gap* and the social welfare losses due to lack of insurance. I prove in Appendix D that the higher the *gap* that opens up at time t the higher the deviation of the time-t allocation from the optimum of a social welfare function that weights the two countries according to their previous wealth. More specifically, the first-best plan is an (infinite) sequence of allocations engineered by a Social Planner who aims at keeping fixed relative countries' wealth from $t = 0$ up to $t = \infty$, weighting countries according to their $t = -1$ relative wealth. If countries' relative wealth has been constant until time $(t-1)$, the *gap* that opens up at time t reflects deviations from this first-best allocation. If countries' relative wealth has changed between $t = -1$ and $(t-1)$ the *gap* that opens up at time t reflects deviations from the second-best plan that would be implemented by a Social Planner who, given the previous deviation from the first-best plan, re-optimizes taking into account the new level of relative wealth. In both cases, the *gap* maps monotonically into social welfare losses due to imperfect risk-sharing.[13]

[13] This holds under the maintained assumption that deviations from the law of one price are contained.

3.2 A new empirical method

Every deviation of the *gap* from zero measures a percentual deviation from full risk-sharing that may result in welfare losses, and is contingent on the shocks that hit the economies at each point in time. This measure of insurance provides the theoretical foundations for a new empirical method to estimate cross-border risk-sharing. The method consists in estimating the *contingent risk-sharing inefficiency* between two countries by using some statistic that quantifies the volatility of the *gap* computed on their Stochastic Discount Factors around zero, within a certain time horizon. The next section shows in detail how this method can be brought to the data, by estimating the deep parameters that characterize the SDFs and by testing the validity in the data of important assumptions. For the moment, we will assume that we know the "true" functional form of the SDFs of the countries under consideration. We will also assume that the data do not reject the equalization of expected SDFs

$$E_t\left(m_{t+1}\right) = E_t\left(m_{t+1}^{\star}\right)$$

Then the unconditional mean of the *gap* (the log difference between SDFs) is equal to zero (see equation (12)). Therefore, the variance and the standard deviation of the *gap* can be used as measures of its volatility around zero, i.e. measures of the contingent risk-sharing inefficiency between the two countries over a certain period of time.

While the variance of the *gap* quantifies the risk-sharing inefficiency, this measure is silent about its causes. It does not indicate which are the sources of idiosyncratic risk against which countries are not fully insured. Moreover a high variance could be due either to the high frequency and strength of uninsurable shocks, or to a low degree of insurance against macro risk. In order to disentangle these effects and to study which sources of macroeconomic risk are left uninsured among two countries, we can regress the time-series for the *gap* on different sources of risk. Assuming there are only productivity disturbances, as in our DSGE model, we can run

$$gap_t = \alpha + \beta \cdot \Delta \log Y_t + \gamma \cdot \Delta \log Y_t^{\star} + \varepsilon_t \tag{13}$$

where Δ denotes first differences. The regressors capture the sources of macroeconomic risk that country H and F should ideally pool. The coefficient associated with each shock quantifies the percentual deviation from full risk-sharing that is created by a 1% variation in the corresponding macro aggregate, thus proxying for the *degree of insurance* against that particular source of risk. The lower the coefficient estimated, say, for country H GDP growth, the higher the degree of insurance against this source of risk. We can assess if some sources of risk are perfectly pooled by testing whether the corresponding coefficient equals zero. We can also estimate (13) over two distinct time periods and decompose changes in the variance of the *gap* over time into changes in the coefficients and changes in the variance of the regressors. This allows to attribute reductions in the risk-sharing inefficiency either to reductions in the extent of macro risk to be pooled (the variance of regressors) or to increases in the degree of insurance against different sources of risk (reductions in the coefficients).[14]

[14] A rise in the degree of insurance against some disturbances could be due to three factors. First, an improvement in insurance possibilities to share risk across countries, e.g. an enlargement in the set of financial instruments traded internationally. Second, some other structural change in a shock process that makes the *available* set of assets more suitable to pool this specific source of risk across countries. For instance, two countries that can trade only in one uncontingent bond can self-insure against idiosyncratic risk quite efficiently if shocks are temporary. If disturbances are permanent, the same asset structure does not allow to insure against any share of risk (see Viani (2010)). Finally, a rise in the degree of insurance

A note on the choice of deflators Combining H and F budget constraints, it is easy to see that the *gap* can be expressed as

$$\begin{aligned}\Delta \log Q_t - \rho\left(\Delta \log C_t - \Delta \log C_t^{\star}\right) &= -\mu\underbrace{(\Delta \log GDP_t - \Delta \log GDP_t^{\star}) + \omega \Delta \log TOT_t}_{\Delta \log v_t} \\ &\quad + \phi \Delta \log NSAV_t - \kappa \Delta \log NFI_t + \Delta \log Q_t\end{aligned}$$

where TOT_t are country H terms of trade, μ, ω, ϕ, and κ are positive constants, and ν_t is the real relative value of the H and the F endowments. Thus, we have two options for estimating equation (13). We can either deflate output using the national CPIs or GDP deflators.[15] If we choose the former, running (13) is equivalent to estimating

$$gap_t = \alpha + \beta \cdot (\Delta \log Y_t + \Delta \log TOT_t) + \gamma \cdot (\Delta \log Y_t^{\star} + \Delta \log TOT_t) + \varepsilon_t$$

The regressors include the impact of terms of trade fluctuations on the relative value of endowments. Thus the coefficients proxy for the degree of insurance against GDP risk, abstracting from the insurance provided by relative price fluctuations that affect the relative value of endowments. Notice that this adjustment is not enough to clear completely the coefficients from the effects of price fluctuations, which enter (13) through real exchange rate movements.

Instead, if we deflate output using its own deflator, running (13) is equivalent to estimating

$$gap_t = \alpha + \beta \cdot (\Delta \log Y_t) + \gamma \cdot (\Delta \log Y_t^{\star}) + \varepsilon_t$$

The coefficients proxy for the degree of insurance provided both by financial markets and relative price fluctuations. In this paper I choose this second solution – using national GDP deflators. The reason is the following. Recent contributions to the open-macro literature (see for instance Corsetti et al. (2008)) emphasize that the structure of financial markets can influence the insurance properties of relative prices. In this sense, the impact on cross-border insurance of financial market integration may encompass also improvements in the insurance properties of relative prices, which should be included in our measure of the degree of insurance.[16]

3.3 The new method in simulations

I apply the new method to the simulated model to check if it identifies correctly full risk-sharing in the two-equity case, and if it captures the correct ranking of insurance among asset structures. Table 4 compares our measure of the bilateral risk-sharing inefficiency (the standard deviation of the *gap*), across different asset structures and values of the trade elasticity. For the Financial autarky case, this statistic signals correctly a high cross-border insurance from relative price fluctuations for $\theta = 1.2$, and a low degree of

could be due to an increase in the insurance properties of relative price fluctuations, either brought about by stronger financial deepness or not.

[15] Notice that, from a purely theoretical point of view, deflating output with GDP deflators is not entirely correct in the other methods (for instance in the approach of Asdrubali, Sorensen and Yosha (1996)). Since these empirical frameworks are based on an identity, using GDP deflators is equivalent to assuming no relative price fluctuations (see section 2).

[16] Simulations carried out in our simple model (available upon request) show that the *difference* between the coefficients estimated using GDP and CPI deflators reveals whether relative price movements reduce or amplify the wealth divergence that would arise at constant prices. The implementation of this exercise if left for future research.

insurance from terms of trade movements for $\theta = 0.3$. Thus, contrary to the coefficients estimated through some of the standard methods, the standard deviation of the *gap* gives an indication of the welfare losses due to the absence of financial markets. Moreover, the standard deviation of the *gap* reflects the correct ranking of insurance among different asset trade regimes, and leads to a non-rejection of full insurance in the complete markets case.

Table 4: Gap standard deviation

Gap standard deviation	$\theta = 0.3$	$\theta = 0.6$	$\theta = 1.2$
Financial autarky	12,89	4,51	2,24
Bond	0,09	0,03	0,02
Complete markets	0	0	0

Table 5 shows the OLS coefficients estimated by running (13). Since the model is fully symmetric $\beta = -\gamma$. Standard errors are negligible and are not reported. The magnitude of the coefficients reflects the degree of insurance against national shocks (the lower the coefficient, the higher the degree of insurance), their sign indicates which country is made relative richer/poorer by the disturbances. Namely $\beta < 0$ (> 0) indicates that an increase in the growth rate of Home GDP makes Foreign consumers relatively poorer (richer) (see the interpretation of the *gap* outlined above).

Table 5: Deviation from full risk-sharing due to
a 1 percent increase in H GDP growth rate

β=$-\gamma$	$\theta = 0.3$	$\theta = 0.6$	$\theta = 1.2$
Financial autarky	-6,61	2,28	-1,11
Bond	-0,03	0,011	-0,005
Complete markets	0	0	0

4 Empirical strategy

4.1 Estimating Stochastic Discount Factors

The only assumptions we need to make to implement our method are the existence of a representative agent in each country, and the specific form of the SDFs. In this paper I adopt two strategies to estimate SDFs. First, I rely on the assumptions of standard open-macro models and postulate a constant risk-aversion; I estimate from the data the deep parameters characterizing the SDFs, and verify that the functional form assumed is not rejected by the data. Second, given the uncertainty on the specific form of the SDFs, I follow Campbell and Cochrane assuming a utility function with time-varying risk-aversion; I adopt the parametrization that was used in this work to replicate the equity-premium puzzle and to capture the history of US stock prices.

4.1.1 Stochastic Discount Factors from the traditional open-macro model

In accordance with what is assumed by most models of international interdependence, I assume that for all countries the period utility function is separable in consumption and depends on the current consumption level according to a standard CRRA function. Namely

$$U(C^i,.) = \frac{(C_t)^{1-\rho}}{1-\rho} + \kappa \tag{14}$$

where ρ is the (common) coefficient of relative risk-aversion and the utility function is allowed to depend on other variables (κ) that do not enter the marginal utility of consumption. (14) implies that the *gap* between country i and j arising at time t can be computed as

$$\begin{aligned} gap_t^{i,j} &\equiv \log m_t^i - \log m_t^j = \\ &= \Delta \log Q_t^{i,j} - \rho\left(\Delta \log C_t^i - \Delta \log C_t^j\right) + \log \beta^i - \log \beta^j \end{aligned} \tag{15}$$

where β^i and β^j denote the discount factors. In accordance with the literature, the deep parameters appearing in this formula can be estimated from the data by testing the equality of countries' expected SDFs. The following equation can then be tested through non-linear GMM

$$\begin{aligned} E_t\left(m_{t+1}^i\right) &= E_t\left(m_{t+1}^j\right) \Rightarrow \\ \frac{\beta^i}{\beta^j} \cdot E_t\left\{\left(\frac{C_{t+1}^i}{C_t^i}\right)^{-\rho} \frac{P_{t+1}^i}{P_t^i}\right\} &= E_t\left\{\left(\frac{C_{t+1}^j}{C_t^j}\right)^{-\rho} \frac{P_{t+1}^j}{P_t^j} \frac{S_{t+1}^{i,j}}{S_t^{i,j}}\right\} \end{aligned} \tag{16}$$

where $S^{i,j}$ is the bilateral nominal exchange rate. This estimation allows to pin down ρ and the relative discount factor $\left(\frac{\beta^i}{\beta^j}\right)$. Moreover, if (16) is not rejected, the mean of the *gap* cannot be rejected to be zero. Therefore, the variance and the standard deviation of $gap^{i,j}$ can be conveniently used as measures of its volatility around zero, i.e. measures of the contingent risk-sharing inefficiency over a period of time.[17] Notice also that failure to reject equation (16) implies that the data do not reject our assumption about the specific form of SDFs.

[17] In the application of this methodology to the US and OECD countries that is described in the next section, equation (16) was not rejected. Clearly, if (16) was rejected we could still use the *second raw moment* of the *gap* to quantify its volatility around zero over a certain period of time, but the validity of the estimated ρ would be questionable.

4.1.2 Stochastic Discount Factors with time-varying risk-aversion

The second strategy consists in assuming a functional form for SDFs following the literature on habit formation, in particular Campbell and Cochrane (1999). The period utility function depends on a stock of habit, and relative risk-aversion is time-varying and state-dependent. Campbell and Cochrane (1999) show that a standard model featuring this SDF solves the short and long-run equity premium puzzle, and captures much of the history of US stock prices. Rabitsch (2008) shows that the same preferences give a rationale for deviations from uncovered interest parity and match the volatility of exchange rates observed in the data. Following these contributions I assume that the utility function is

$$\frac{(C_t - X_t)^{1-\rho} - 1}{1-\rho} \tag{17}$$

where X_t is the habit level. The relationship between consumption and habit is captured by the surplus consumption ratio

$$S_t = \frac{C_t - X_t}{C_t}$$

Thus relative risk aversion (the local curvature of the utility function) is state-dependent and related to S_t by

$$\eta_t = -\frac{U_{CC} C}{U_C} = \frac{\rho}{S_t}$$

Assume further

$$\log S_{t+1} = (1-\phi)\log \bar{S} + \phi \log S_t + \lambda(\log S_t)(\log C_{t+1} - \log C_t - \log G)$$

where G is the average growth rate of the economy and

$$\begin{aligned}
\lambda(\log S_t) &= \frac{1}{\bar{S}}\sqrt{1 - 2\left(\log S_t - \log \bar{S}\right)} - 1, \log S_t \leq \log S_{\max} \\
\lambda(\log S_t) &= 0, \log S_t > \log S_{\max}
\end{aligned}$$

and

$$\bar{S} = \rho\sqrt{\frac{\rho}{1-\phi-\frac{B}{\rho}}}$$

The Stochastic Discount Factors of H and F consumers read respectively

$$m_t = \beta\left(\frac{S_{t+1}}{S_t}\right)^{-\rho}\left(\frac{C_{t+1}}{C_t}\right)^{-\rho}\frac{P_t}{P_{t+1}}$$

$$m_t^{\star} = \beta^{\star}\left(\frac{S_{t+1}^{\star}}{S_t^{\star}}\right)^{-\rho}\left(\frac{C_{t+1}^{\star}}{C_t^{\star}}\right)^{-\rho}\frac{P_t^{\star}}{P_{t+1}^{\star}}\frac{\varepsilon_t}{\varepsilon_{t+1}}$$

The *gap* is given by

$$\begin{aligned}
gap_t \equiv{} & \log(m_t) - \log(m_t^{\star}) = \\
& \Delta \log Q_t - \rho\left(\Delta \log C_t - \Delta \log C_t^{\star} + \Delta \log S_t - \Delta \log S_t^{\star}\right)
\end{aligned} \tag{18}$$

For each country I compute G as the average growth rate of national consumption from the data. Following Campbell and Cochrane (1999) and Rabitsch (2008), I assume $\rho = 2$, $\phi = 0.985$, and $B = -0.01$.

4.2 Estimating the risk-sharing inefficiency

Having checked for both specifications of SDFs that the mean of the *gap* cannot be rejected to be zero in the data, I use its variance and standard deviation as measures of its volatility around zero. The *contingent inefficiency* between country i and j over the period $t = 1, ..., T$ will be measured by [18]

$$\hat{\sigma}^2\left(gap^{i,j}\right) = \sum_{t=1}^{T}\left[\left(gap^{i,j}\right) - E\left(gap^{i,j}\right)\right]^2 \qquad \text{or} \qquad \hat{\sigma}\left(gap^{i,j}\right) = \left[\hat{\sigma}^2\left(gap^{i,j}\right)\right]^{1/2}$$

4.3 Detecting significant changes in the inefficiency over time

Significant variations in the risk-sharing inefficiency between country i and j can be found by testing for structural changes in the variance of $gap^{i,j}$. Since the dates of potential changes are unknown, I find it convenient to employ the Quasi-Maximum Likelihood methodology of Qu and Perron (2007), which allows to detect structural breaks at unknown dates, find their number, and test for their significance. More precisely, I estimate

$$gap_t^{i,j} = \mu + \varepsilon_t \tag{19}$$

where μ is a constant and ε_t are *i.i.d.* Normal disturbances. I assume no break occurred in μ, and test for structural breaks in the variance-covariance matrix of errors.[19]

4.4 Macroeconomic risk versus degree of insurance

Following the same reasoning as for equation (13), if the relationship between country i and j was described by a model with exogenous GDP, government spending and monetary shocks, plus disturbances coming from currency speculation and from productivity in the rest of the world, it is easy to show that the risk-sharing inefficiency between country i and j would follow – in a first order approximation to the model – the following process

$$\begin{aligned} gap_t^{i,j} \quad &= \quad \alpha \cdot \Delta \log gdp_t^i + \beta \cdot \Delta \log gdp_t^j + \gamma \cdot \Delta \log gov_t^i \\ &\quad +\delta \cdot \Delta \log gov_t^j + \eta \cdot \Delta \log s_t^{i,j} + \zeta \cdot \Delta \log gdp_t^{world} + \varepsilon_t \end{aligned} \tag{20}$$

where $\Delta \log gdp_t^i$, $\Delta \log gdp_t^j$, $\Delta \log gdp_t^{world}$ denote respectively the growth rate of country i, country j and "rest of the world" GDP, deflated through national GDP deflators and not including government expenditure.[20] Rest of the world GDP affects the inefficiency between country i and j as long as it has an asymmetric impact on consumption in the two countries.[21] $\Delta \log gov_t^i$, $\Delta \log gov_t^j$, and $\Delta \log s_t^{i,j}$ are country i and country j final government expenditure expressed in local currency, and the nominal exchange rate between country i and country j currency. Nominal exchange rate growth is capturing both the effects of monetary shocks and the impact of possible disturbances coming from currency speculation. I use OECD GDP growth as a proxy for rest of the world GDP.

[18] I will use the standard deviation to compare the extent of the inefficiency across countries, and the variance to detect changes in the inefficiency over time and to investigate their causes.

[19] See Qu and Perron (2007) for additional details on the tests.

[20] Equation (20) could also include lags of the independent variables as regressors.

[21] For instance, if a shock to Italian GDP had an asymmetric impact on the US and the UK, it would contribute to determine $gap^{US,UK}$. Including rest of the world GDP as a regressor captures this effect, reducing the potential omitted-variable bias. Clearly $\Delta \log gdp^{world}$ does not include country i and country j's GDP.

I estimate equation (20) instrumenting nominal exchange rate growth with GDP and exchange rate growth lagged values.[22]

Two kinds of analyses can be performed using (20). First we can establish if some sources of risk are perfectly pooled by testing whether the corresponding coefficients equal zero. Second, we can run (20) over two distinct time horizons in order to attribute changes in the variance of the *gap* over time to changes in the coefficients (the degree of insurance against shocks) and changes in the variance of regressors (the extent of macroeconomic risk to be pooled), according to the following procedure. First, any reduction in the variance of $gap^{i,j}$ between the two subperiods *pre* and *post* (denoting the pre and post-break years) can be expressed in terms of changes in the variance of shocks and changes in the coefficients in (20) as

$$\begin{aligned}
&Var\left(gap^{i,j}\right)_{post} - Var\left(gap^{i,j}\right)_{pre} = \\
&\left(\hat{\alpha}^2_{post} - \hat{\alpha}^2_{pre}\right) \cdot Var\left(\Delta \log\left(gdp^i\right)\right)_{pre} + \\
&\hat{\alpha}^2_{pre} \cdot \Big(Var\left(\Delta \log\left(gdp^i\right)\right)_{post} - Var\left(\Delta \log\left(gdp^i\right)\right)_{pre}\Big) + \\
&\left(\hat{\alpha}^2_{post} - \hat{\alpha}^2_{pre}\right) \cdot \Big(Var\left(\Delta \log\left(gdp^i\right)\right)_{post} - Var\left(\Delta \log\left(gdp^i\right)\right)_{pre}\Big) + ...
\end{aligned} \tag{21}$$

where the subscripts denote the pre and post-break years, and $\hat{\alpha}_{pre}$ and $\hat{\alpha}_{post}$ are the coefficient on country i GDP estimated on the two subsamples. Second, based on the decomposition in (21) we can compute the fraction of the reduction in $Var\left(gap^{i,j}\right)$ due to changes in each coefficient and in the variance of each shock. For instance, the share of the reduction in the variance of $gap^{i,j}$ across the two periods that is due to changes in α is given by

$$share\,(\alpha) = \frac{V\,(\alpha)}{Var\,(gap^{i,j})_{post} - Var\,(gap^{i,j})_{pre}} \tag{22}$$

where $V\,(\alpha)$ is the reduction in $Var\left(gap^{i,j}\right)$ explained by a fall in α (see Appendix E for details on its computation). $share\,(\alpha)$ can be interpreted as the share of the reduction in the contingent inefficiency due to better insurance against country i GDP risk.

Analogously we can compute the fraction of the change in $Var\left(gap^{i,j}\right)$ due to a fall in the variance of each shock, that is in the extent of risk to be shared. For instance for country i GDP

$$share\left(gdp^i\right) = \frac{V\left(gdp^i\right)}{Var\,(gap^{i,j})_{post} - Var\,(gap^{i,j})_{pre}} \tag{23}$$

where $V\left(gdp^i\right)$ is the reduction in the variance of the *gap* explained by a fall in the volatility of country i GDP across the two subsamples.

5 Risk-sharing among the US and OECD countries

The new methodology is applied to estimate the bilateral risk-sharing inefficiency between the US and industrialized OECD countries and to study its evolution over time.

[22] Lags 1-3 or 3 only depending on the SDFs specification. I use Instrumental Variables as there could be other macro aggregates omitted in the regression that affect both the *gap* and the regressors. This problem is most likely to be relevant for a relative price, the nominal exchange rate, whose volatility is larger in the data and (partially) unexplained by fundamentals. Also a constant, not reported in the tables, was included in the regression.

Sample and data The US is the reference country, the partner countries are 16 OECD industrialized economies: 4 extra-EU countries (Australia, Canada, Japan, and New Zealand), 8 Euro-Area countries (Belgium, Finland, France, Ireland, Italy, the Netherlands, Portugal, and Spain), and 4 European countries that did not join the Euro (Norway, Sweden, Switzerland, and the UK).[23] The data are sampled at the quarterly frequency. Data on final consumption expenditure, consumption deflators, nominal exchange rates, government final consumption expenditure, and GDP, are from the OECD Economic Outlook. Data on population at the annual frequency were obtained from the OECD National Accounts. The period under consideration is 1970:2-2008:2.

5.1 SDFs from the traditional model: nominal exchange rate volatility as a source of risk

Deep parameters estimation Assume CRRA utility function according to (14). Following Dedola and De Fiore (2005), I tested equation (16) simultaneously for all partner countries vis-a-vis the US through non-linear GMM.[24] I used as instruments for each equation a constant, lags 1 and 2 of consumption factor growth of the US and partner country j, US and country j inflation, and nominal depreciation of the US dollar.

Table 10 in Appendix F shows the results. Although in the specification of the SDFs we do not allow for taste shocks, the model seems to capture well the data, consistently with the results of previous GMM estimations of the same equation.[25] The p-value of the J statistic indicates that the model cannot be rejected. The relative discount factors of the US and country j, $\left(\frac{\beta^{US}}{\beta^{j}}\right)$ cannot be rejected to be equal to 1 for all partner countries. The common coefficient of relative risk-aversion is estimated to be 1.32, well in the range of estimates derived in previous studies.[26]

[23] Some OECD-members were not included in the sample because of data availability.

[24] I use continously-updated GMM and employ the Newey-West adjustment for heteroskedasticity and autocorrelation. Given the interdependence between countries, it is more efficient to test the equations for all partners simultaneously. Testing each equation individually, though, did not lead to reject the model, consistently with the findings of Kollmann (1995).

The coefficient of risk-aversion is assumed to be the same for all countries. I found that, when assuming a different ρ for each partner, the model cannot be rejected, but some of the coefficients are estimated very imprecisely and are not significant. The coefficients that are significant are all between zero and 2. Therefore, since the main goals of the GMM estimation implemented here are (1) to establish whether the model can be rejected and (2) to estimate risk-aversion, this coefficient is assumed to be the same for all countries, which leads (again) not to reject the model and yields significant estimates for all the parameters. It should be noted that the common ρ estimated this way is in the range of the significant ones estimated assuming different risk-aversion coefficients for each country.

[25] See, for instance, the results of Kollmann (1995) for several pairs of OECD economies, and Dedola and De Fiore (2005) for European countries vis-a-vis Germany. These results are robust to the inclusion of one additional instrument lag. I have chosen a reduced set of instruments in order to minimise the potential bias that could stem from the excess of overidentifying restrictions in small samples.

[26] Stock and Wright (2000) show that when esimating risk-aversion *inside a country* (by using the traditional CAPM Euler Equation and national panel data), the estimated coefficient may be very low because of a weak instrument problem. It should be noticed, however, that *cross-country estimations* of the same equation tend to find low values for the the CRRA, but seem to be immune from the weak-instrument issue. Using non-linear GMM estimation and a cross-country version of the traditional CAPM Euler Equation, Dedola and De Fiore (2005) estimate a common CRRA of 0.39 for Euro countries. Selaive and Tuesta (2003) estimate the common CRRA of Australia and the US to be 1.58. By using the method of Stock and Wright (2000) to check for the presence of weak instruments, they establish that their result is not driven by this potential problem. Corsetti and Dedola (2006) also test the cross-country version of the traditional Euler equation through non-linear GMM and find it cannot be rejected for many countries vis-a-vis the US. They argue that testing the cross-country version of this equation may mitigate problems due to measurement errors and to the presence of potential peso problems *common* to the two countries. This might be the reason why the standard CAPM model is sometimes rejected in other studies that test its one-country version.

Although the risk-aversion coefficient is estimated very precisely, the estimation might not be immune from measurement errors in consumption and weak instrument issues. Since this could bias downwards the estimation, robustness checks on all results have been performed for higher values of this coefficient (up to $\rho = 7$) and are available upon request.[27]

Contingent risk-sharing inefficiency and its decrease over time Given these results, for every partner country j in the sample I compute $gap_t^{US,j}$ according to equation (15), exploiting the equality of the estimated discount factors for the US and each partner country ($\beta^{US} = \beta^j, \forall j$). Since the GMM estimation of equation (16) did not reject the equality of expected SDFs we can safely use the standard deviation (and the variance) of $gap^{US,j}$ as a measure of the contingent risk-sharing inefficiency between the US and country j.

Table 11 shows the standard deviation of $gap^{US,j}$ for each partner country j, using the ρ estimated from the data and other higher values. The estimated standard deviation of the *gap* between the US and Canada (in the first row) is statistically smaller than the standard deviation of the *gap* between the US and any other partner country.[28] Results are robust to higher values of the risk-aversion coefficient. Over the period 1970-2008 the smallest contingent risk-sharing inefficiency has been the one between the US and Canada. Notice that this result could be due either to strong financial linkages between the two countries (a high degree of insurance via asset trade) or to a significant synchronization in their business cycles that could reduce the extent of purely idiosyncratic risk that the two countries should pool.

Figures 1 and 2 graph the *gap* between the US and, respectively, the UK and the Netherlands over the last forty years. The volatility of the *gap* between the US and these countries seems to decrease from the mid-Nineties. This intuition is confirmed by the results of the structural break analysis to detect significant changes in the variance of the *gap*, reported in Table 12. For every partner country, I show the number of breaks detected through the QML tests of Qu and Perron (2007), their dates, and the percentual reduction in the inefficiency across the pre- and post-break period. Asterisks next to the number of breaks indicate their significance according to the test of no change versus a number k of breaks.[29]

The only breaks in the variance that are found to be statistically significant and robust to all the values ρ are those estimated in 1992-1993 for several European countries, Italy,

[27] See Carroll (2001) for the downward bias due to measurement errors in consumption, and Stock and Wright (2000) for weak instrument issues.

[28] I use a Wald test to establish whether this difference is statistically significant. The test statistic is

$$W = [r(b) - q]' \left\{ R(b)\hat{V}R'(b) \right\}^{-1} [r(b) - q] \sim \chi^2(1) \tag{24}$$

where b is a vector of variances and covariances needed for the specific statistic, and $r(b)$ is a function that maps these variances into the statistic of interest, in this case

$$\left[\hat{\sigma}\left(gap^{US,CAN}\right) - \hat{\sigma}\left(gap^{US,j}\right) \right], j \neq CAN$$

and $R(b) = \partial r(b) / \partial b'$. $\hat{V}$ is the variance-covariance matrix of b and is estimated employing the Newey-West adjustment for autocorrelation and heteroskedasticity with a lag length of $T^{1/4}$, where T is the sample size. Results are reported in Table 3.4 where asterisks next to the standard deviation of each $gap^{US,j}$ indicate it is statistically different from $\hat{\sigma}\left(gap^{US,CAN}\right)$. As usual, an asterisk indicates significance at the 10% level, two at the 5% level, and three at the 1% level.

[29] See Qu and Perron (2007). Additional details on the estimation– including the estimated variance of errors before and after the breaks, and results for intermediate values of the risk-aversion coefficient – are available upon request. When computing the percentual reduction in the inefficiency, the pre-break period is assumed to start in 1978 in order to get subsamples of similar length.

the Netherlands, Spain, Sweden, Switzerland, UK, Portugal, and Belgium.[30] From the early Nineties the risk-sharing inefficiency between the US and these countries fell by 30 to 60%, depending on the partner. Figure 3 shows the rolling-window estimated standard deviation of the *gap* and its trend for some selected partners, illustrating this fall in the risk-sharing inefficiency.[31]

Macroeconomic risk vs degree of insurance I estimate equation (20) for the pre and post-break periods (1978-1992 and 1993-2008) for the countries involved in the 1992-1993 break. Results are shown in Tables 13 and 14. Recall that each coefficient quantifies the percentual deviation from full risk-sharing associated with a 1% variation in the corresponding macro aggregate, and is therefore an inverse proxy for the degree of insurance against that particular source of risk. Asterisks next to the coefficients denote their significance computed using Newey-West adjusted standard errors with maximum lag length 10. $(\cdot)$ symbols next to 1993-2008 coefficients indicate that the coefficient changed significantly over time, and show the level of significance.[32] The vast majority of coefficients are found to be significant, signaling that perfect insurance against the main sources of macro risk can be rejected in most of the cases. Only for some of the partner countries there has been a significant improvement in the degree of insurance against some sources of risk.

Tables 15 and 16 report the shares of the reduction in the variance of the *gap* between the pre and post-break period explained by falls in the variance-coviariance of shocks and by changes in the coefficients. Shares were computed according to (22) and (23) using only statistically significant coefficients at the 10% level and variations in coefficients over time significant at the 10% level. Only for 3 countries out of 8 (the Netherlands, Spain, and the UK) more than 20% of the reduction in the welfare losses from imperfect risk-sharing can be attributed to improvements in the degree of insurance. For the other countries involved in the 1992-1993 break, most of the reduction in the risk-sharing inefficiency vis-a-vis the US was due to a decrease in nominal exchange rate volatility.[33] Thus constant risk-aversion implies that nominal exchange rate volatility does not foster an efficient allocation of resources across countries. It is rather an important source of macroeconomic risk whose decrease led to a significant reduction in insurance inefficiencies.

5.2 SDFs with time-varying risk-aversion: the effects of the Great Moderation

Contingent risk-sharing inefficiency and its decrease over time Assume the utility function with habits in (17). I adopt the parametrization of Campbell and Cochrane (1999). Table 17 shows the standard deviation of $gap^{US,j}$ for each partner country j. For each partner the mean of the *gap* could not be rejected to be zero in the data.

Table 19 shows the results of the structural break analysis on the variance of the *gap*. The risk-sharing inefficiency is found to decrease significantly over time for all partner countries but Ireland, Norway, and Portugal, for which significant increases were detected. The breaks seem to be clustered around two periods: last-Seventies/beginning-Eighties and end of the Nineties. Their magnitude varies between 20 and 60%. Figure 4 shows the

[30] A second break was detected for Italy, the Netherlands and Spain at the beginning of the Eighties, but it was found not to be robust to slightly higher values of the risk-aversion coefficient.

[31] The years indicated in the axes of the graphs correspond to the last year in each window. The series are smoothed by HP-filtering.

[32] As in Dedola and De Fiore (2005), the null $\hat{\alpha}_{pre} = \hat{\alpha}_{post}$ was tested separately on the outcomes of the pre and post-break regressions. The post-break coefficient is statistically different from the pre-break one if and only if the null of equality is rejected by both tests.

[33] Also changes in the covariance of the regressors contribute to explain the reduction in the variance of the *gap*. I report them in the tables only when their effect is non-negligible.

rolling-window estimated standard deviation of the *gap* and its trend for some selected partners. The inefficiency between the US and, respectively, Switzerland and Belgium was subject to two reductions, at the beginning of the Eighties and at the end of the Nineties. The inefficiency relative to Italy and the Netherlands fell significantly only at the end of the Nineties.

Macroeconomic risk vs degree of insurance For the partner countries for which a reduction in the inefficiency was detected, I estimate (20) for the pre- and post-break period using as regressors also lagged values of the independent variables.[34] Results are shown in Tables 20 – 23. Perfect insurance against the main sources of macro risk can be rejected in most of the cases. Half of the breaks are characterized by generalized falls in the coefficients, while the evidence is mixed for the remaining half.

Tables 24 and 25 report the shares of the reduction in the variance of the *gap* explained by reductions in the volatility of shocks and by changes in the coefficients. Half of the reductions in insurance inefficiencies (the UK from 1984, Switzerland from 1983 and 1999, the Netherlands from 1996, Finland from 1978, and Italy from 1998) were mostly due to a higher degree of insurance against macro shocks. The other half of the reductions were due to a fall in the volatility of macro aggregates, mainly GDP of both the US and the partner countries, and its components. This is true for the breaks detected for Japan in 1978 and 1997, Canada in 1997, the UK in 1999, and Belgium in 1983. Notice that the beginning of a structural reduction in the volatility of US GDP (the beginning of the so-called Great Moderation) has been usually set at 1984 by the empirical literature on business-cycle fluctuations. This is consistent with the reduction in the risk-sharing inefficiency detected between the US and Belgium around 1983, which is found to be largely driven by falls in the volatility of US GDP and government spending. Summers (2005) finds that the Great Moderation began around the mid-Seventies in Japan, and documents another drop in the volatility of Japanese GDP around 1997-1998. Structural break analysis identifies a significant reduction in the variance of Canadian government spending in 1997 (results are available upon request). The timing of these episodes of reductions in macro volatility is consistent with the break-dates we estimate.

Thus time-varying risk-aversion implies that a reduction in the extent of risk to be shared internationally, requires a decrease in the volatility of macro fundamentals, as the one that occurred for several countries during the Great Moderation years.

6 Concluding remarks

This paper contributes to the literature on empirical risk-sharing measurement by investigating in simulations the robustness of existing empirical methods to the presence of international price fluctuations. I also propose a new methodology to quantify international insurance, which is valid also in presence of relative price movements, and apply it to study the evolution of risk-sharing among industrialized countries. The main results have been outlined in the main text and will not be repeated here. This section will instead discuss some additional applications of this method, left for future research.

The empirical analysis carried out assuming constant risk-aversion unveiled that the risk-sharing inefficiency between the US and some European economies fell from the early

[34] Using lagged values of the independent variables as regressors improves significantly the results. This is not surprising given that the risk-aversion coefficients used to compute SDFs and *gap* are history-dependent and thus embed the effects of past shocks. The inclusion of lagged regressors was not necessary in the case of constant risk-aversion.

In a few cases independent variables could explain only a small share of the variance of the *gap*. These regressions are not reported.

Nineties, mostly due to a decrease in nominal exchange rate volatility. The years 1992-1993, in which the break in the contingent inefficiency is detected, coincide with the signing of the Maastricht Treaty and with the end of the Exchange Rate Mechanism (ERM) crisis in Europe. Some of the partner countries for which the break was detected were directly involved in the ERM crisis (namely, Sweden, the UK, and Italy). All partners involved in the break (except Switzerland) signed the Maastricht Treaty, thus bounding themselves to take part in the process of European integration and possibly to join the common currency. The end of currency turmoils combined with strong expectations of exchange rate stabilization might have been the factors driving the decrease in nominal exchange rate volatility between the US and European countries. Whether the same events reduced risk-sharing inefficiencies within European countries, and whether this results would be robust to other SDFs specifications, is an interesting issue on which the method proposed in this paper could shed light.

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(Chapter head:)

A Other methods

A.1 Obstfeld (1993)

Given any two countries, assume representative agents, no taste shocks, and CRRA utility function. Complete markets imply

$$\frac{\beta^t (C_t)^{-\rho}}{(C_0)^{-\rho}} = \frac{\beta^{\star t} (C_t^\star)^{-\rho}}{(C_0^\star)^{-\rho}}$$

In a two-country environment, the method of Obstfeld (1993) consists in estimating

$$\log(C_t) = \log(C_t^\star) + \log(C_0/C_0^\star) + \log(\beta/\beta^\star)(t/\rho)$$

Assuming equal discount factors (as in our model) and differencing, we can estimate[35]

$$\Delta\log(C_t) = \alpha + \beta\cdot\Delta\log(C_t^\star) + \gamma\cdot\Delta\log(Y_t) + \varepsilon_t \tag{25}$$

Markets are complete if $\beta = 1$ and $\gamma = 0$.

Using the simulated model described in section 2, I test equation (25) deflating national output using local CPIs. The results are reported in Table 6. They are qualitatively similar to those derived using the method of Asdrubali, Sorensen and Yosha (1996), and unveil that Obstfeld's framework is subject to the same issues due to the presence of relative price fluctuations.

A.2 Brandt, Cochrane and Santa Clara (2006)

Given any two countries, Brandt, Cochrane and Santa Clara (2006) propose the following index of international risk-sharing

$$\mu \equiv 1 - \frac{\sigma^2\left(\log\frac{RER_{t+1}}{RER_t}\right)}{\sigma^2(\log m_{t+1}) + \sigma^2\left(\log m_{t+1}^\star\right)}$$

where

$$m_{t+1} = -\rho(\log\Delta C_{t+1})$$

$$m_{t+1}^\star = -\rho\left(\log\Delta C_{t+1}^\star\right)$$

$\mu = 1$ is interpreted as perfect risk-sharing.

Table 7 reports the results from applying this method to our model. Full insurance is rejected even in the model specification in which asset markets are effectively complete. The reason is the following. According to this methodology the variance of Stochastic Discount Factors expressed in local currencies (m and $m^\star$) represents the extent of risk to be shared, and every deviation from Purchasing Power Parity is considered a manifestation of the risk-sharing inefficiency. Viani (2010) shows however that an efficient cross-border allocation can be characterized only by relative consumption in conjuncture with relative prices. In this sense, from a theoretical point of view not all price fluctuations reflect a departure from full risk-sharing.

[35] See Obstfeld (1993) for the need to carry out an estimation using first-differenced data.

Table 6: OLS estimation of β and γ in Obstfeld regression

β, γ	$\theta = 0.3$	$\theta = 0.6$
Financial autarky	$\beta = 0$, $\gamma = 1$	$\beta = 0$, $\gamma = 1$
Bond	$\beta = -0,27$, $\gamma = 1,27$	$\beta = -0,086$, $\gamma = 1,086$
Complete markets	$\beta = -0,273$, $\gamma = 1,273$	$\beta = -0,087$, $\gamma = 1,087$

β, γ	$\theta = 1.2$
Financial autarky	$\beta = 0$, $\gamma = 1$
Bond	$\beta = 0,158$, $\gamma = 0,841$
Complete markets	$\beta = 0,159$, $\gamma = 0,84$

Table 7: Risk-sharing index of Brandt, Cochrane and Santa Clara

μ (risk-sharing index)	$\theta = 0.3$	$\theta = 0.6$	$\theta = 1.2$
Financial autarky	-0,13	-2,74	0,89
Bond	0,12	0,27	0,5
Complete markets	0,09	0,28	0,5

A.3 Flood, Marion and Matsumoto (2008)

Flood, Marion and Matsumoto (2008) measure the risk-sharing inefficiency using the variance of a country's consumption on world consumption. In our setup this translates to

$$\lambda = var\left(\log\left(C_t\right) - \log\left(C_t + C_t^{\star}\right)\right)$$

$\lambda = 0$ is interpreted as perfect risk-sharing.

Table 8 shows the results from applying this method to our simulated model. Clearly, the benchmark of efficiency on which the methodology is based does not take into account relative price fluctuations, leading both to a rejection of full risk-sharing in the complete markets case and to problems in capturing the correct ranking of insurance across asset structures.

Table 8: Risk-sharing index of Flood, Marion and Matsumoto

λ (inefficiency index)	$\theta = 0.3$	$\theta = 0.6$	$\theta = 1.2$
Financial autarky	2	0,04	0,39
Bond	0,61	0,33	0,15
Complete markets	0,5	0,28	0,14

B A note on the use of GDP deflators

I apply the ASY method to our simulated model without implementing any correction for the presence of relative price fluctuations. Namely, I deflate output using national GDP deflators. Results are shown in the following Table. In this case, the ASY method is subject only to two of the three issues described in the main text.

The coefficients estimated signal a higher degree of insurance under Financial Autarky than under complete markets for $\theta = 0.6$ because the relative value of endowments does not reflect asymmetric wealth effects. Full insurance is always rejected in the two-equity model because the benchmark of risk-sharing on which the method is based does not consider that the efficient consumption allocation depends also on relative prices.

On the other hand, the method captures the correct ranking of insurance across asset trade regimes for $\theta = 0.3$. Contrary to the implementation that uses national CPIs to deflate output, the present version of the method is not affected by the third, typical issue due to relative price fluctuations – the fact that efficient financial markets may channel resources "upstream". This happens because the coefficients estimated in this version capture also the effects of relative price fluctuations on the relative value of endowments.

Table 9: ASY method with GDP-deflated output
OLS estimation of the fraction of shocks left uninsured

β (fraction uninsured)	$\theta = 0.3$	$\theta = 0.6$	$\theta = 1.2$
Financial autarky	1,76	0,29	0,85
Bond	0,938	0,744	0,534
Complete markets	0,934	0,746	0,532

C Insurance measure and changes in relative wealth

Using the model outlined in section 2, I show that the *gap* reflects changes in countries' relative wealth, for any asset structure we may assume in the model.

C.1 Changes in relative wealth under Financial autarky

First, let's consider the Financial autarky case. Households have no means of smoothing consumption over time and in each period their consumption must equal the value of their income. H and F budget constraints read

$$C_t P_t = P_{ht} Y_{ht} \qquad C_t^\star P_t^\star = P_{ft} Y_{ft}$$

Agents' intra-temporal decision can be solved through a standard expenditure minimization.[36] The inter-temporal problem of H agents can be solved through a standard Lagrangian

$$\max_{\{C_t, \chi_t\}} . : L = \frac{(C_t)^{1-\rho}}{1-\rho} + \chi_t \left(P_{ht} Y_{ht} - P_t C_t \right), \forall t$$

where χ_t is the multiplier attached to H agents' budget constraints, and represents the shadow value of current income, $P_{ht} Y_t$. The first order condition with respect to consumption reads

$$(C_t)^{-\rho} = \chi_t P_t \tag{26}$$

Its analogue for the F agent is given by

$$(C_t^\star)^{-\rho} = \chi_t^\star P_t^\star \tag{27}$$

where ${\chi_t}^\star$ is the shadow value of F household's current income.

Taking the ratio of (26) and (27), and dividing it by its analogue at time $(t-1)$ gives

$$\left(\frac{C_t^\star}{C_t} \right)^{-\rho} \left(\frac{C_{t-1}^\star}{C_{t-1}} \right)^{\rho} = \frac{Q_t}{Q_{t-1}} \left(\frac{\chi_t}{\chi_t^\star} \right) \left(\frac{\chi_{t-1}^\star}{\chi_{t-1}} \right)$$

Taking logs

[36] See for instance Obstfeld and Rogoff (1996).

$$\Delta \log (Q_t) - \rho (\Delta \log (C_t) - \Delta \log (C_t^\star)) = gap_t = [\Delta \log (\chi_t) - \Delta \log (\chi_t^\star)]$$

where Δ denotes the first difference of a variable. The *gap* arising at time t in response to country-specific shocks reflects asymmetric changes of the shadow value of income in the two countries.[37]

C.2 Changes in relative wealth with trade in uncontingent bonds

Assume trade in uncontingent bonds. This amounts to modifying H household's budget constraint to

$$P_t^B B_{t+1} = P_{ht} Y_{ht} - P_t C_t + P_t B_t$$

where B_{t+1} denotes bonds purchased in period t, P_t^B their unitary price in terms of the H consumption bundle. Bonds are assumed to be in zero net supply. H households' inter-temporal problem is given by

$$\max_{\{C_t, B_{t+1}, \nu_t\}} . : L = E_0 \left\{ \sum_{t=0}^{\infty} \beta^t \left[\frac{(C_t)^{1-\rho}}{1-\rho} + \nu_t \left(P_{ht} Y_{ht} + P_t B_t - C_t P_t - P_t^B P_t B_{t+1} \right) \right] \right\}$$

v_t represents the shadow value of income at time t. It can be shown that time t income equals the present discount value of wealth

$$(P_{ht} Y_t + B_t P_t - P_t C_t) = \sum_{j=1}^{\infty} \prod_{i=0}^{j-1} P_{t+j}^B \left[P_{t+j} C_{t+j} - P_{h,t+j} Y_{t+j} \right] \equiv PDV \ wealth$$

Hence, the multiplier ν_t is the shadow value of wealth at time t.

H agent's first order conditions are given by

$$(C_t)^{-\rho} = \nu_t P_t \tag{28}$$

$$P_t^B = \beta E_t \left\{ \frac{\nu_{t+1}}{\nu_t} \frac{P_{t+1}}{P_t} \right\} \tag{29}$$

From H agent's first order condition with respect to consumption and from its analogue for the F consumer, we get

$$\Delta \log (Q_t) - \rho (\Delta \log (C_t) - \Delta \log (C_t^\star)) = gap_t = [\Delta \log (\nu_t) - \Delta \log (\nu_t^\star)] \tag{30}$$

The *gap* arising at time t corresponds to asymmetric changes in the shadow value of wealth in the two countries.[38]

[37] If H and F agents' discount factors (the $\beta' s$) were different, the gap would reflect asymmetric wealth effects *plus* the (log) difference between the discount factors β and $\beta^\star$

$$gap_t = [\Delta \log (\chi_t) - \Delta \log (\chi_t^\star)] + [\log (\beta) - \log (\beta^\star)]$$

This shows that if agents are heterogeneous in the way they discount the future, for risk to be perfectly shared their relative wealth should change at the rate $[\log (\beta) - \log (\beta^\star)]$.

[38] If $\beta \neq \beta^\star$, equation (30) would read

$$gap_t = [\Delta \log (\nu_t) - \Delta \log (\nu_t^\star)] + [\log (\beta) - \log (\beta^\star)]$$

The proof above refers to a setup in which consumers can only smooth *ex-post* their consumption. When agents can trade in contingent assets, gross portfolio holdings can be used to hedge *ex-ante* against idiosyncratic risk: in response to any disturbance assets' return differentials deliver automatic transfers of wealth into consumers' budget constraints. Viani (2010) proves that if households can form an optimal portfolio that hedges them ex-ante – to some extent – against shocks, the *gap* still represents asymmetric changes in the shadow value of wealth in the two countries – the *gap* arising at time t is still described by equation (30). In this case, the asymmetric wealth effects mirrored in $[\Delta \log(\nu_t) - \Delta \log(\nu_t^\star)]$ are *residual* after portfolio returns have delivered contingent transfers of income in households' budget constraints.

D Insurance measure and welfare

The first-best plan that delivers the maximum social welfare attainable in the workhorse model can be characterized as the outcome of a Social Planner maximization. The Social Planner must allocate consumption between the H and the F country, and that weights symmetrically the two economies. She solves the following problem

$$\max.: E_0 \left\{ \sum_{t=0}^{\infty} \Omega \beta^t \frac{(C_t)^{1-\rho}}{1-\rho} + (1-\Omega)\,\beta^t \frac{(C_t^\star)^{1-\rho}}{1-\rho} \right\} \tag{31}$$

$$s.t.: C_t = \left[(\delta)^{1/\theta} (C_{ht})^{\frac{\theta-1}{\theta}} + (1-\delta)^{1/\theta} (C_{ft})^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{\theta-1}} \tag{32}$$

$$C_t^\star = \left[(\delta)^{1/\theta} \left(C_{ft}^\star\right)^{\frac{\theta-1}{\theta}} + (1-\delta)^{1/\theta} \left(C_{ht}^\star\right)^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{\theta-1}} \tag{33}$$

$$Y_t = C_{ht} + C_{ht}^\star \tag{34}$$

$$Y_t^\star = C_{ft} + C_{ft}^\star \tag{35}$$

In what follows μ_t and $\mu_t^\star$ denote the multipliers attached to the CES bundle constraints, that is the shadow value of consumption, respectively, in the H and in the F country. In the decentralized problem whose solution corresponds to the Social Planner allocation, the ratio $\left(\frac{\mu_t^\star}{\mu_t}\right)$ corresponds to the real exchange rate Q_t. The first order conditions with respect to C_t and $C_t^\star$ read

$$\Omega \cdot (C_t)^{-\rho} = \mu_t \tag{36}$$

$$(1-\Omega) \cdot (C_t^\star)^{-\rho} = \mu_t^\star \tag{37}$$

The sequence of equations (36) and (37) evaluated at $t = 0, ...\infty$ characterize the allocation that constitutes the first-best plan that (by construction) delivers the maximum social welfare attainable in this economy, that is the sequence of allocations that maximize the Social Welfare function in (31). Equations (36) and (37) characterize the time t allocation that is part of this plan. The first-best allocations may coincide with the ones attained in a decentralized model with complete asset markets. In particular, in order for the Planner's solution to correspond to a complete-markets decentralized allocation, the Planner's weights Ω and $(1-\Omega)$ should represent countries' relative wealth determined by the initial conditions C_{-1}, $C^{\star}_{-1}$, μ_{-1}, and $\mu^{\star}_{-1}$, that is relative wealth at time $t = -1$.[39] From (26), (27), and (28) we know that relative wealth at time t is pinned down by the following relationship

$$\left(\frac{\omega_t}{\omega^{\star}_t}\right) = \left(\frac{C_t}{C^{\star}_t}\right)^{-\rho} \frac{P^{\star}_t}{P_t} \tag{38}$$

where ω_t and $\omega^{\star}_t$ denote the shadow value of wealth of the H and the F consumer at time t.[40] Therefore, the ratio of the Social Planner's weights must satisfy

$$\frac{\Omega}{1-\Omega} = \left(\frac{\omega_{-1}}{\omega^{\star}_{-1}}\right) = \left(\frac{C_{-1}}{C^{\star}_{-1}}\right)^{-\rho} \frac{P^{\star}_{-1}}{P_{-1}}$$

Since only the ratio of the weights matters for the optimization, I assume without loss of generality

$$\Omega = \frac{(C_{-1})^{-\rho}}{P_{-1}} \qquad \text{and} \qquad (1-\Omega) = \frac{\left(C^{\star}_{-1}\right)^{-\rho}}{P^{\star}_{-1}}$$

Assume that countries' relative wealth has always been constant from $t = -1$ until time $(t-1)$ - i.e. assume the realized allocations have always coincided with the original first-best plan, that is conditions (36) and (37) have never been violated up to time $(t-1)$. Then

$$\Omega = \frac{(C_{-1})^{-\rho}}{P_{-1}} = \frac{(C_0)^{-\rho}}{P_0} = ... = \frac{(C_{t-1})^{-\rho}}{P_{t-1}}$$

$$(1-\Omega) = \frac{\left(C^{\star}_{-1}\right)^{-\rho}}{P^{\star}_{-1}} = \frac{(C^{\star}_0)^{-\rho}}{P^{\star}_0} = ... = \frac{\left(C^{\star}_{t-1}\right)^{-\rho}}{P^{\star}_{t-1}}$$

Substituting these expressions into the first order conditions (36) and (37) yields the benchmark allocation at time t expressed as a function of the previous period allocation.

$$\left(\frac{C_t}{C_{-1}}\right)^{-\rho} = \frac{\mu_t}{\mu_{-1}} \tag{39}$$

$$\left(\frac{C^{\star}_t}{C^{\star}_{-1}}\right)^{-\rho} = \frac{\mu^{\star}_t}{\mu^{\star}_{-1}} \tag{40}$$

In economies in which asset markets are incomplete the Planner's first order conditions (36) and (37) need not hold with equality. Assume that the incomplete market allocations have always coincided with the first-best plan from $t = 0$ up to time $(t-1)$, but at time t an idiosyncratic shock makes the incomplete market allocation deviate from the benchmark allocation. In this case

$$\left(\frac{C^{IM}_t}{C^{IM}_{t-1}}\right)^{-\rho} \cdot \varphi_t = \frac{\mu^{IM}_t}{\mu^{IM}_{t-1}} \tag{41}$$

$$\left(\frac{C^{\star IM}_t}{C^{\star IM}_{t-1}}\right)^{-\rho} \cdot \varphi^{\star}_t = \frac{\mu^{\star IM}_t}{\mu^{\star IM}_{t-1}} \tag{42}$$

[39] See Ljungqvist and Sargent (2004).

[40] The shadow value of wealth in (38) may equal the shadow value of current income or the shadow value of the present discounted value of lifetime resources, depending on the asset structure assumed. Equation (38), however, holds for every asset structure - even in a setup with multiple assets.

where φ and $\varphi^{\star}$ represent the wedge between marginal utility of consumption and CPI, and the IM superscript denotes the incomplete markets allocation. Taking logs of (41) and (42) gives

$$\log\left(\varphi_t^{\star}\right) = \log\left(\mu_t^{\star IM}\right) + \rho\log\left(C_t^{\star IM}\right) - \log\left(\mu_{t-1}^{\star IM}\right) - \rho\log\left(C_{t-1}^{\star IM}\right)$$

$$\log\left(\varphi_t\right) = \log\left(\mu_t^{IM}\right) + \rho\log\left(C_t^{IM}\right) - \log\left(\mu_{t-1}^{IM}\right) - \rho\log\left(C_{t-1}^{IM}\right)$$

Notice that if the shock is purely idiosyncratic, $\log\left(\varphi_t\right)$ and $\log\left(\varphi_t^{\star}\right)$ must have opposite signs, that is if $\log\left(\varphi_t\right)$ is positive, $\log\left(\varphi_t^{\star}\right)$ must be negative and viceversa.[41] The incomplete markets *gap* coincides with the difference between $\log\left(\varphi_t^{\star}\right)$ and $\log\left(\varphi_t\right)$

$$gap_t = \Delta\log\left(Q_t^{IM}\right) - \rho\left(\Delta\log\left(C_t^{IM}\right) - \Delta\log\left(C_t^{\star IM}\right)\right) = \log\left(\varphi_t^{\star}\right) - \log\left(\varphi_t\right)$$

The opposite sign of $\log\left(\varphi_t^{\star}\right)$ and $\log\left(\varphi_t\right)$ is sufficient to ensure that a higher *gap* (in absolute value) must be generated by an incomplete markets allocation that implies higher wedges between marginal utility of consumption and CPI, therefore a higher deviation from the Social Planner's (logged) optimality conditions (36) and (37). Due to the concavity of the Social Welfare function in (31) as a sum of concave functions, a higher deviation from the Social Planner's focs maps into a lower social welfare. Therefore, the higher the *gap* that arises in response to shocks in any incomplete markets setup, the higher the loss in social welfare caused by the disturbance with respect to the full risk-sharing Planner's benchmark.

If countries' relative wealth has changed between $t = -1$ and time $(t-1)$ the *gap* that opens up at time t reflects deviations from the second-best plan that would be implemented by a Social Planner who, given the previous deviation from the first-best plan, re-optimizes taking into account the new level of relative wealth.

Assume without loss of generality that only one deviation from the first-best plan has occurred, and that it took place at time $(\tau-1) < (t-1)$. The second-best Planner's problem reads

$$\max.: E_{\tau}\left\{\sum_{j=0}^{\infty}\Omega_{\tau-1}\beta^j\frac{\left(C_{\tau+j}\right)^{1-\rho}}{1-\rho} + \left(1-\Omega_{\tau-1}\right)\beta^j\frac{\left(C_{\tau+j}^{\star}\right)^{1-\rho}}{1-\rho}\right\} \tag{43}$$

subject to constraints (32)-(35) and to the initial conditions $C_{\tau-1}$, $C_{\tau-1}^{\star}$, $\mu_{\tau-1}$, $\mu_{\tau-1}^{\star}$. The subscript $(\tau-1)$ attached to the weights signals that relative weights are chosen so as to reflect the new ranking of wealth determined by the time τ shock. Following the same steps as above

$$\Omega_{\tau-1} = \frac{\left(C_{\tau-1}\right)^{-\rho}}{P_{\tau-1}} = \frac{\left(C_{\tau}\right)^{-\rho}}{P_{\tau}} = ... = \frac{\left(C_{t-1}\right)^{-\rho}}{P_{t-1}}$$

$$\left(1-\Omega_{\tau-1}\right) = \frac{\left(C_{\tau-1}^{\star}\right)^{-\rho}}{P_{\tau-1}^{\star}} = \frac{\left(C_{\tau}^{\star}\right)^{-\rho}}{P_{\tau}^{\star}} = ... = \frac{\left(C_{t-1}^{\star}\right)^{-\rho}}{P_{t-1}^{\star}}$$

These characterize the benchmark time t allocation that delivers the maximum social welfare attainable in this economy given initial conditions $C_{\tau-1}$, $C_{\tau-1}^{\star}$, $\mu_{\tau-1}$, $\mu_{\tau-1}^{\star}$, i.e. the time t component of the second-best sequence of allocations that maximizes the Social

[41] If this was not the case one of the resource constraints would be violated.

Welfare function in (43). Following the same steps as above it's easy to show that the higher the *gap* arising in any incomplete markets setup at time t, the higher the deviation from focs (39) and (40), that is the higher the social welfare loss due to imperfect risk-sharing.

E Decomposing changes in the variance of the gap

The reduction in the variance of $gap^{i,j}$ explained by a fall in the coefficient α, associated with country i GDP growth, is computed as

$$
\begin{aligned}
V(\alpha) = {} & (\hat{\alpha}_{post} - \hat{\alpha}_{pre}) Var\left(\Delta \log gdp^i\right)_{pre} + \\
& \frac{1}{2}(\hat{\alpha}_{post} - \hat{\alpha}_{pre}) \left(Var\left(\Delta \log gdp^i\right)_{post} - Var\left(\Delta \log gdp^i\right)_{pre}\right) + \\
& (\hat{\alpha}_{post} - \hat{\alpha}_{pre}) \sum_{\lambda=\beta,\gamma,\delta,\eta,\zeta} 2\hat{\lambda}^{pre} Cov\left(\Delta \log gdp^i, L^\lambda\right)_{pre} + \\
& (\hat{\alpha}_{post} - \hat{\alpha}_{pre}) \sum_{\lambda=\beta,\gamma,\delta,\eta,\zeta} \hat{\lambda}^{pre} \left(Cov\left(\Delta \log gdp^i, L^\lambda\right)_{post} - Cov\left(\Delta \log gdp^i, L^\lambda\right)_{pre}\right) + \\
& (\hat{\alpha}_{post} - \hat{\alpha}_{pre}) \sum_{\lambda=\beta,\gamma,\delta,\eta,\zeta} \left(\hat{\lambda}^{post} - \hat{\lambda}^{pre}\right) Cov\left(\Delta \log gdp^i, L^\lambda\right)_{pre} + \\
& (\hat{\alpha}_{post} - \hat{\alpha}_{pre}) \sum_{\lambda=\beta,\gamma,\delta,\eta,\zeta} \left(\hat{\lambda}^{post} - \hat{\lambda}^{pre}\right) \frac{2}{3} \left(Cov \Delta \log\left(gdp^i\right), L^\lambda_{post} - Cov\left(\Delta \log gdp^i, L^\lambda\right)_{pre}\right)
\end{aligned}
$$

where $\lambda = \beta, \gamma, \delta, \eta, \zeta$ denote all other coefficients in equation (20) and L^λ the source of shock with which each coefficient λ is associated. The subscripts pre and post attached to second moments denote the subperiod (pre- or post-break) for which they are computed. The cross-product terms between the variation in the coefficient and the change in the variance are typically assumed to be negligible for small variations.

F Figures and Tables

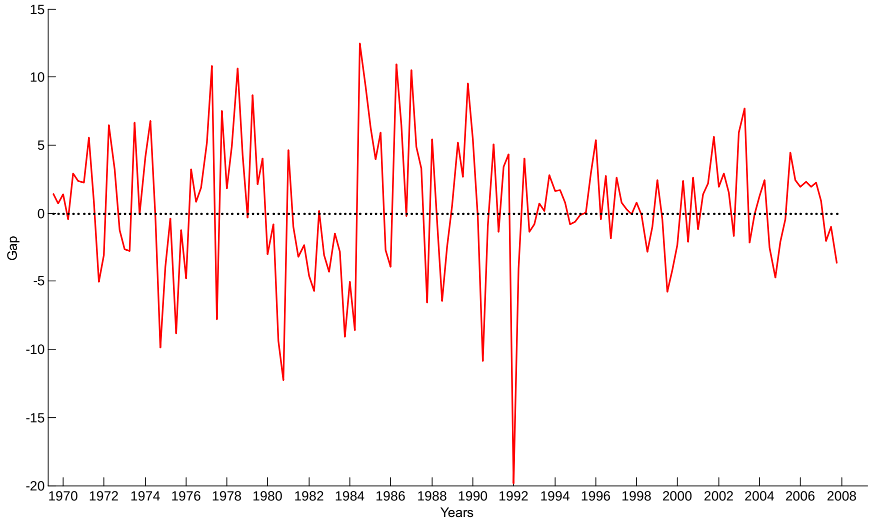

Figure 1 – Percentual deviations from full risk-sharing (*gap*) between the US and the UK

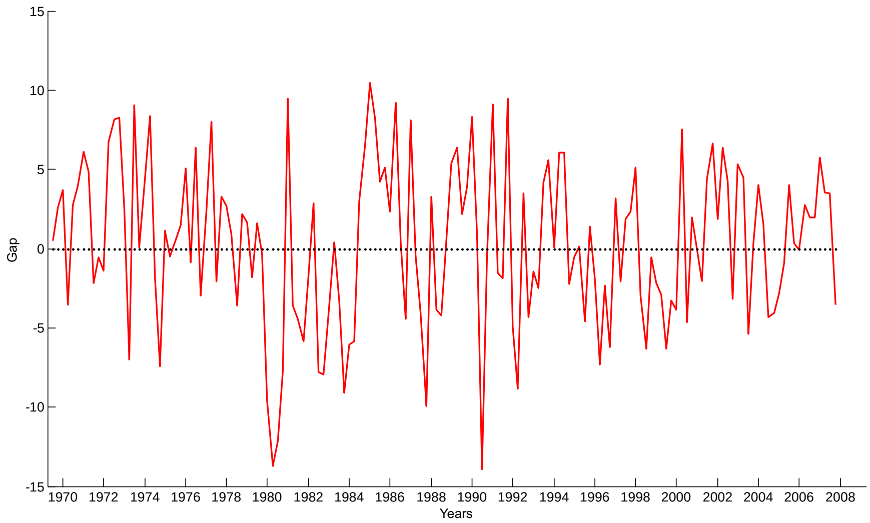

Figure 2 – Percentual deviations from full risk-sharing (*gap*) between the US and the Netherlands

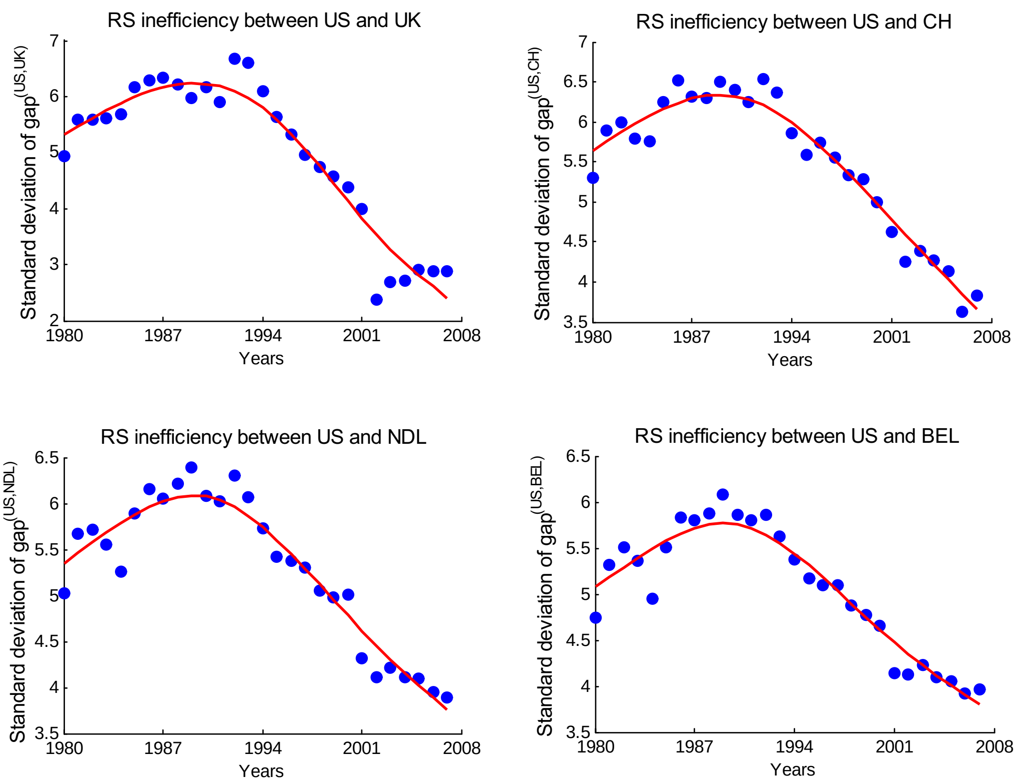

Figure 3 – 10-years rolling-window estimates of the standard deviation of the *gap* for selected partner countries (CRRA preferences). The series are smoothed by HP-filtering.

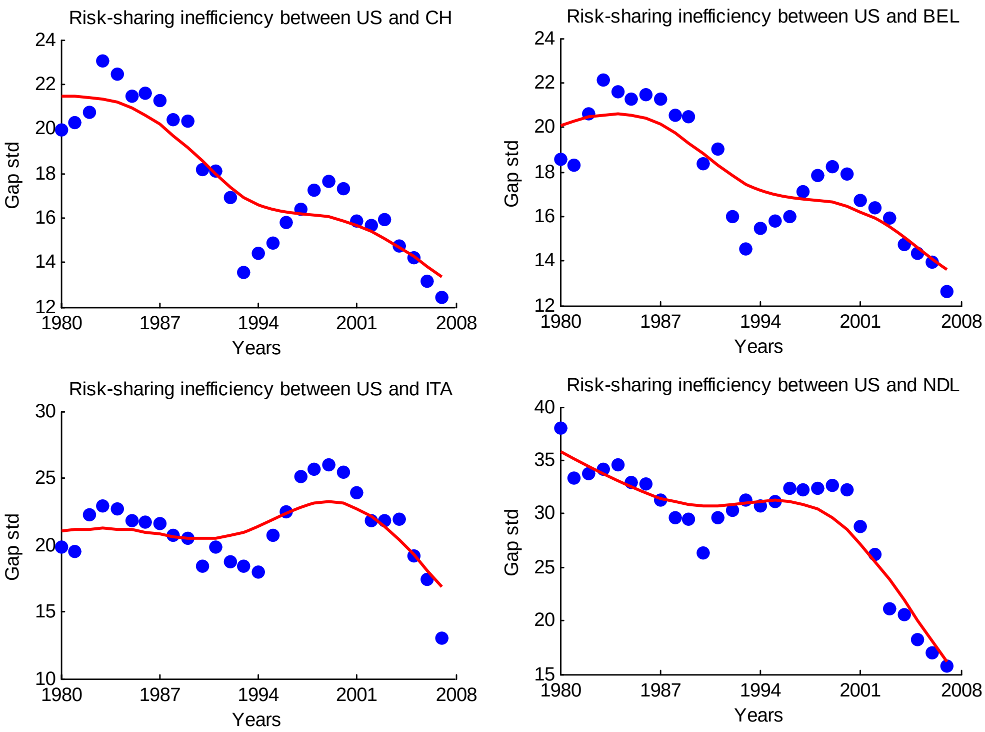

Figure 4 – 10-years rolling-window estimates of the standard deviation of the *gap* for selected partner countries (CC preferences). The series are smoothed by HP-filtering.

Table 10: Results from GMM estimation

	Parameter	Standard error	P-value	J-stat	P-value
Risk-aversion coefficient	1,32	0,0676	6,25E-085	1,709	1
Rel discount factor AUS	1,0025	0	0,0000		
Rel discount factor CAN	1,0004	0,0009	0,0000		
Rel discount factor FRA	1,0043	0,0015	0,0000		
Rel discount factor ITA	1,0047	0,0017	0,0000		
Rel discount factor JAP	1,0064	0,0015	0,0000		
Rel discount factor NDL	1,0068	0,0017	0,0000		
Rel discount factor ESP	1,0053	0,0016	0,0000		
Rel discount factor SWE	1,0048	0,0020	0,0000		
Rel discount factor CH	1,0102	0,0015	0,0000		
Rel discount factor UK	1,0026	0,0016	0,0000		
Rel discount factor IRL	1,0024	0,0015	0,0000		
Rel discount factor NOR	1,0021	0,0017	0,0000		
Rel discount factor NZL	1,0058	0,0016	0,0000		
Rel discount factor POR	1,0040	0,0017	0,0000		
Rel discount factor BEL	1,0045	0,0016	0,0000		
Rel discount factor FIN	1,0032	0,0018	0,0000		

NOTE: Continuously-updated GMM. Newey-West adjustment for heteroskedasticity and autocorrelation.

Table 11: Contingent risk-sharing inefficiency between the US and the partner countries (1970-2008) as measured by the std of the gap (CRRA preferences)

Partner country	$\rho = 1.32$	$\rho = 4$	$\rho = 5$
CAN	2,5	4,39	5,21
AUS	4,18 ***	5,65 ***	6,39 **
FR	4,70 ***	5,84 ***	6,41 **
ITA	4,63 ***	5,99 ***	6,70 **
JAP	5,24 ***	7,05 ***	7,98 ***
NDL	5,05 ***	6,97 ***	7,89 ***
ES	4,89 ***	6,17 ***	6,80 ***
SWE	5,03 ***	7,81 ***	9,09 ***
CH	5,23 ***	6,05 ***	6,51 **
UK	4,77 ***	6,70 ***	7,64 ***
IRL	4,57 ***	6,07 ***	6,86 ***
NOR	4,53 ***	7,02 ***	8,22 ***
NZL	5,29 ***	9,05 ***	10,75 ***
POR	5,07 ***	7,65 ***	8,81 ***
BEL	4,87 ***	6,05 ***	6,66 ***
FIN	4,9 ***	7,30 ***	8,47 ***

NOTE: Asterisks next to each standard deviation denote the level of significance (* denotes the 10% level, ** the 5%, *** the 1%) of the Wald test for the null hypothesis $\sigma(gap^{US,CAN}) = \sigma(gap^{US,j})$.

Table 12: Reduction in the risk-sharing inefficiency between the US and the partner countries, detected as structural breaks in the variance of the gap (CRRA preferences)

Partner country	No. breaks	Date	% reduction
ITA	2 ***	1993	36,74%
NDL	2 ***	1993	35,15%
ES	2 ***	1993	35,15%
SWE	1 ***	1993	34,39%
CH	1 ***	1993	38,34%
UK	2 ***	1992	60,31%
POR	1 ***	1993	35,19%
BEL	1 *	1993	32,13%
AUS	0	/	/
CAN	0	/	/
FR	0	/	/
JAP	0	/	/
IRL	0	/	/
NOR	0	/	/
NZL	0	/	/
FIN	0	/	/

NOTE: Asterisks next to the number of breaks indicate their significance (* for 10%, ** for 5%, *** for 1%) according to the test of no change versus a number k of breaks. See Qu and Perron (2007) for details on this test.

Table 13: IV regressions – dependent variable: gap(US,j), CRRA preferences

UK	1978-1992	1993-2008	**ITA**	1978-1992	1993-2008
US gdp	-0,97 ***	-0,63 *** (·)	US gdp	-1,06 ***	-0,52 *** (·)
UK gdp	1,02 ***	0,05 (···)	ITA gdp	0,81 **	0,94 ***
US govt	-0,97 ***	-0,6 ***	US govt	-1,11 ***	-0,55 *** (··)
UK govt	1,06 ***	-0,39 (···)	ITA govt	0,61	0,33
NER	1,12 ***	0,99 *** (·)	NER	1,00 ***	1,04 ***
OECD gdp	-0,17	-0,32	OECD gdp	0,11	-0,45 *** (···)
NDL	1978-1992	1993-2008	**ESP**	1978-1992	1993-2008
US gdp	-0,75 ***	-1,19 *** (··)	US gdp	-0,85 ***	-0,42 *** (···)
NDL gdp	0,19 ***	0,49 *** (··)	ESP gdp	0,36 ***	0,63 ***
US govt	-0,52 **	-0,92 *** (·)	US govt	-0,72 ***	-0,38 *** (···)
NDL govt	-0,22	0,2	ESP govt	0,53 ***	0,73 ***
NER	1,11 ***	0,87 *** (··)	NER	1,04 ***	0,98 *** (·)
OECD gdp	0,03	-0,14	OECD gdp	0,03	-0,03

NOTE: * next to the coefficients denotes their significance computed using Newey-West adjusted standard errors with maximum lag length 10. (· next to 1993-2008 coefficients indicates that the coefficient changed significantly with respect to its 1978-1992 value, and shows the level of significance.

Table 14: IV regressions – dependent variable: gap(US,j), CRRA preferences

SWE	1978-1992	1993-2008	**CH**	1978-1992	1993-2008
US gdp	-0,97 ***	-0,49 ** (··)	US gdp	-0,74 ***	-0,62 ***
SWE gdp	0,34 ***	0,34	CH gdp	0,29 *	0,3 ***
US govt	-0,69 **	-0,56 **	US govt	-0,78 ***	-0,54 ***
SWE govt	0,07	0,53 *	CH govt	0,65 ***	0,32 *** (·)
NER	1,01 ***	0,99 ***	NER	1,01 ***	0,98 ***
OECD gdp	0,64 **	-0,12 (··)	OECD gdp	-0,2 *	-0,31 ***

POR	1978-1992	1993-2008	**BEL**	1978-1992	1993-2008
US gdp	-0,63 ***	-0,42 **	US gdp	-0,9 ***	-0,66 ** (··)
POR gdp	-0,2	0,23	BEL gdp	0,6 **	1,13 *** (·)
US govt	-0,66 ***	-0,7 ***	US govt	-0,91 ***	-0,79 ***
POR govt	0,16	0,32	BEL govt	0,97 **	1,1 **
NER	1,03 ***	0,94 ***	NER	0,95 ***	1,02 *** (··)
OECD gdp	-0,02	-0,13	OECD gdp	0,02	-0,29 *** (··)

NOTE: See note to Table 3.6

Table 15: Share of the risk-sharing inefficiency between the US and the partner country explained by falls in the variance of shocks and improvements in the degree of insurance (CRRA preferences)

UK		**ITA**	
US gdp	1,66	US gdp	5,52
UK gdp	-1,92	ITA gdp	0
US govt	0	US govt	-4,54
UK govt	1,3	ITA govt	0
NER	19,23	NER	0
OECD gdp	0	OECD gdp	-0,47
Insurance	**20,27%**	**Insurance**	**0,51%**
Var(US gdp)	1,87	Var(US gdp)	5,5
Var(UK gdp)	2,88	Var(ITA gdp)	-0,18
Var(US gov)	0,25	Var(US gov)	0,79
Var(UK gov)	1,72	Var(ITA gov)	0
Var(NER)	71,26	Var(NER)	83,31
Var(OECD gdp)	0	Var(OECD gdp)	0
Cov(US gdp,NER)	2,46	Cov(US gdp,NER)	0,5
		Cov(ITA gdp,NER)	9,92
Shocks	**80,43%**	**Shocks**	**99,85%**
NDL		**ESP**	
US gdp	-6,16	US gdp	4,6
NDL gdp	-0,41	ESP gdp	0
US govt	2,1	US govt	-2,94
NDL gov	0	ESP gov	0
NER	43,65	NER	20,3
OECD gdp	0	OECD gdp	0
Insurance	**39,18%**	**Insurance**	**21,96%**
Var(US gdp)	1,68	Var(US gdp)	4,24
Var(NDL gdp)	0,14	Var(ESP gdp)	0,87
Var(US gov)	0,11	Var(US gov)	0,4
Var(NDL gov)	0	Var(ESP gov)	2
Var(NER)	57,62	Var(NER)	74,78
Var(OECD gdp)	0	Var(OECD gdp)	0
Shocks	**59,54%**	**Shocks**	**82,29%**

NOTE: Shares are computed using statistically significant coefficients at the 10% level

Table 16: Share of the risk-sharing inefficiency between the US and the partner country explained by falls in the variance of shocks and improvements in the degree of insurance (CRRA preferences)

SWE		**CH**	
US gdp	3,54	US gdp	0
SWE gdp	0	CH gdp	0
US govt	0	US govt	0
SWE gov	0	CH gov	0,27
NER	0	NER	0
OECD gdp	1,21	OECD gdp	0
Insurance	**4,75%**	**Insurance**	**0,27%**
Var(US gdp)	5,48	Var(US gdp)	2,11
Var(SWE gdp)	0,88	Var(CH gdp)	0
Var(US gov)	0,36	Var(US gov)	0,31
Var(SWE gov)	0	Var(CH gov)	0,32
Var(S)	84,33	Var(S)	96,08
Var(OECD gdp)	0,34	Var(OECD gdp)	0,02
Cov(SWE gdp,NER)	7,64		
Shocks	**99,04%**	**Shocks**	**98,84%**
POR		**BEL**	
US gdp	0	US gdp	5,34
POR gdp	0	BEL gdp	-6,86
US govt	0	US govt	0
POR gov	0	BEL gov	0
NER	0	NER	-34,19
OECD gdp	0	OECD gdp	-0,31
Insurance	**0%**	**Insurance**	**-36,02%**
Var(US gdp)	2,64	Var(US gdp)	6,34
Var(POR gdp)	0	Var(BEL gdp)	0,85
Var(US gov)	0,38	Var(US gov)	0,85
Var(POR gov)	0	Var(BEL gov)	1,3
Var(NER)	95,17	Var(NER)	113,3
Var(OECD gdp)	0	Var(OECD gdp)	0
		Cov(BEL gdp,NER)	11,43
Shocks	**98,18%**	**Shocks**	**134,06%**

NOTE: See note to Table 3.8

Table 17: Contingent risk-sharing inefficiency between the US and the partner countries (1970-2008) as measured by the std of the gap (CC preferences)

Partner country	CC preferences
CAN	23,2
AUS	25,32
FR	21,23
ITA	19,99
JAP	30,25
NDL	32,04
ES	18,26
SWE	41
CH	18,65
UK	27,61
IRL	23,61
NOR	37,04
NZL	41
POR	23,29
BEL	19,16
FIN	36,86

Table 18: Gap mean – Campbell and Cochrane preferences

Partner country	Mean of gap$^{(US,j)}$
AUS	0,08
CAN	-0,09
FRA	0,55
ITA	0,48
JAP	0,12
NDL	-0,01
ES	0,36
SWE	-0,68
CH	0,71
UK	0,25
IRL	0,63
NOR	0,21
NZL	-0,99
POR	0,37
BEL	0,96
FIN	0,04

NOTE: For each partner and CRRA coefficient, the null hypothesis E(gap(US,j))=0 was tested using Newey-West adjustd standard errors with a maximum autocorrelation lag of 10. In no case could the null be rejected.

Table 19: Reductions in the risk-sharing inefficiency between the US and the partner countries, detected as structural breaks in the variance of the gap (CC preferences)

Partner	No. breaks	Date and % reduction	Date and % reduction
AUS	1 ***	1987 47%	/
CAN	2 ***	1977 43%	1997 44%
FRA	2 ***	1978 29%	1997 45%
ITA	1 ***	1998 51%	/
JAP	2 ***	1978 40%	1997 18%
NDL	1 ***	1996 30%	/
ESP	1 ***	1983 28%	/
SWE	2 ***	1978 55%	1998 44%
CH	2 ***	1983 29%	1999 33%
UK	2 ***	1984 39%	1999 43%
IRL	2 ***	1983 44%	2000 -105%
NOR	2 ***	1977 -111%	1999 47%
NZL	2 ***	1987 51%	1997 62%
POR	2 ***	1983 27%	1991 -96%
BEL	2 ***	1983 34%	1999 34%
FIN	1 ***	1978 61%	/

NOTE: See note to Table 3.5

Table 20: IV regressions – dependent variable: gap(US,j), CC preferences

CAN	Break 1		Break 2	
	1970-1977	1977-1984	1986-1997	1997-2008
US gdp	-13,6 ***	-5,81 ** (· · ·)	-12,5 **	-10,9 ***
L1	21,4 ***	3,35 ** (· · ·)	0,8	5,1 **
L2	-8,14 **	-2,53 ** (··)	-0,7	6,1 **
CAN gdp	6,91 **	24,8 *** (· · ·)	16,65 **	6,5 *
L1	-13,03 **	-32,27 *** (· · ·)	-6,2	-3,6
L2	-6,84 ***	25,32 *** (· · ·)	-1,9	-1,2
US govt	-0,05	-2,97	-13 **	-2,7 (··)
L1	1,34	8,64 ** (··)	2,5	2,5
L2	-16,02 ***	-0,61	5,5	1,9
CAN govt	8,72 ***	16,8 ***	21,2 *	1 (·)
L1	-14,3 ***	-32,4 *** (· · ·)	-10	-4
L2	0,08	9,1 ** (··)	-0,1	-2,6
NER	3,71	-6,91 ***	-1,3	2,5 ***
OECD gdp	0,96	-19,4 *** (· · ·)	6,7	-6,5 *

JAP	Break 1		Break 2	
	1970-1977	1977-1984	1986-1997	1997-2008
US gdp	-6,5 **	-5	-8,1 ***	-0,8 (· · ·)
L1	6,5 **	12,6 *** (··)	11 ***	1,6 (· · ·)
L2	-1,2	3,1 ** (··)	-2,9	6,2 ** (··)
JAP gdp	10,6 ***	25,6 *** (· · ·)	13,9 ***	7,7 *** (· · ·)
L1	-11,3 ***	-14 ***	-19,2 ***	-3 *** (· · ·)
L2	1,6	-4,2	6,1 ***	-4,8 *
US govt	-10,9 **	-0,5 (··)	-7,5 ***	-1,3 (· · ·)
L1	10,6	17,9 *** (· · ·)	10,8 ***	1,7 (· · ·)
L2	-11,6 *	13,7 ** (· · ·)	2,7	0,1
JAP govt	17,1 ***	10 *** (··)	2,3	3,7
L1	-18,3 ***	-21 ***	-14,48 ***	-4,4 *** (· · ·)
L2	3,8 ***	-7,6	-0,6	2,5
NER	0,7	0,1	1,1	1,6 ** (··)
OECD gdp	6,3	-9,9	-1,6	-16,6 *** (· · ·)

NOTE: See note to Table 3.6. L1 and L2 denote first and second lags

Table 21: IV regressions – dependent variable: gap(US,j), CC preferences

CH	Break 1		Break 2	
	1970-1983	1983-1996	1990-1999	1999-2008
US gdp	-4,6 *	-11,9 *** (···)	-14,1 ***	-5 *** (···)
L1	13,5 ***	12 ***	12,2 ***	5,7 *** (··)
L2	0,6	-3,5	2,6	0,4
CH gdp	-19,6 **	0,4 (···)	1,1	-1,8
L1	25,6 ***	4,6 *** (···)	-0,2	7,1 *** (···)
L2	-9,1	-2,6	1,12	-2,6 ** (··)
US govt	-11,3 ***	.10,7 ***	-18,4 ***	-6,8 *** (···)
L1	11,7 **	8 ***	12,4 ***	1,6 (···)
L2	8,6	2	8,6 ***	2 (···)
CH govt	-9,6	1,7	11,2 ***	-1,1 (···)
L1	0,1	1,5 * (·)	-0,5	7 ** (··)
L2	9,4 *	-1,6 (·)	-1,6	-1,4
NER	-1,9	0,5 ** (··)	0,6	1,6 ** (··)
OECD gdp	-1,6	1,4	3,6 *	-5,2 ***

UK	Break 1		Break 2	
	1970-1983	1983-1996	1990-1999	1999-2008
US gdp	-8,6 ***	-11 ***	-7,6 **	-5,9 *
L1	11,3 ***	17 *** (·)	10,5 ***	-0,1 (···)
L2	-3,5	-8,2 *** (···)	-4	-9,5 *** (···)
UK gdp	9,1 ***	1,9 (···)	0,6	-1,07
L1	-11 ***	-0,3 (···)	-15,7 **	1,3 (··)
L2	-0,1	-4,4	19,1 ***	10,8 *
US govt	-13,5 *	-9,3 ***	-8 ***	-10,4 ***
L1	3,3	9,5 *** (···)	7,2 **	-0,3 (··)
L2	-14,5 **	-2,1 (··)	0,1	-5,7 *** (··)
UK govt	7,9 **	-0,03 (··)	-7,8	-13,1 ** (··)
L1	-11,2 **	-1,2 (··)	-7,1	7,2
L2	0,3	-7,7	12,9	9,2
NER	1,8 ***	1	-0,9	1,44 ** (··)
OECD gdp	1,2	2,2	7,4	4,4

NOTE: See note to Table 3.13

Table 22: IV regressions – dependent variable: gap(US,j), CC preferences

FRA	1971-1978	1978-1985	**NDL**	1985-1996	1996-2007
US gdp	-17,55 ***	-11,1 ***	US gdp	1,3	-10,7 *** (···)
L1	13,4 ***	9,9 *** (···)	L1	4,4	11 *** (···)
L2	-1,9	1,4	L2	-7,7 ***	6,7 **
FRA gdp	-15,5 **	23,5 ***	NDL gdp	7,1 *	5,1 *
L1	-2,1	14,7 *** (···)	L1	-9,5 ***	-12,8 ***
L2	-5,3 ***	-9,8	L2	2,5	-0,5
US govt	-23,5 ***	-14,3 ***	US govt	3,3	-9,2 *** (···)
L1	15,6 ***	3,5 (···)	L1	-5,6	11 *** (···)
L2	-6,4 *	10,1 ***	L2	0,2	4,7
FRA govt	-33,3 ***	17 (···)	NDL govt	0,7	1,8
L1	-9,8 **	-13,8 (··)	L1	-14,6 **	-7 ***
L2	17,2 ***	-15,4 *	L2	-8,9	-1,8
NER	-1,05	2,3 ***	NER	4,2 **	0,3 (··)
OECD gdp	13,9	-10 ***	OECD gdp	2,4	-0,3

ITA	1987-1998	1998-2008	**FIN**	1970-1977	1977-1984
US gdp	-18,8 ***	-10,8 ***	US gdp	-17,8 **	-16,9 ***
L1	10,5 **	1,05 (··)	L1	25,3 ***	9,9 *** (···)
L2	8,1 *	4,8	L2	29,5 ***	1,2 (···)
ITA gdp	17,4	9,5 ** (··)	FIN gdp	10,8 ***	-12,4 ***
L1	-9,2	-1,8	L1	-26,9 ***	-10,7 *** (···)
L2	3,4	-5,9 *** (···)	L2	2,8	6,2 *** (···)
US govt	-15,5 ***	-7,6 ***	US govt	124,5 ***	-10,5 ** (···)
L1	-2,4	1,7	L1	50,8 ***	13,13 *** (···)
L2	5,7	-2,4	L2	34,9 ***	13,92 ***
ITA govt	34,7 ***	1,7 (···)	FIN govt	-22,2 ***	4,8 (···)
L1	-2,6	7,2 ** (··)	L1	-25,8 ***	-6 (···)
L2	-12,5 *	-7,7 **	L2	71,4 ***	-28,6 *
NER	-2,4 ***	2,5 ***	NER	11,7	0,5
OECD gdp	6,5	-2,5	OECD gdp	29,9 ***	41,1 ***

NOTE: See note to Table 3.13

Table 23: IV regressions – dependent variable: gap(US,j), CC preferences

NZL	1986-1997	1997-2008
US gdp	-5,2	-8,5 *** (· · ·)
L1	27,2 ***	9,5 *** (· · ·)
L2	-17,7 **	-10 ***
NZL gdp	8,7 ***	3,4 (· · ·)
L1	-10,8 ***	-5,7 *** (··)
L2	1,4	0,3
US govt	6,9	-8,2 *** (· · ·)
L1	14,5 **	6,4 **
L2	-5	9,2 *** (· · ·)
NZL govt	4,1 **	3,3 (··)
L1	-12 ***	-7,7 *** (··)
L2	4,7 **	1,3 (··)
NER	5,2 ***	0,9 * (··)
OECD gdp	-13,8 ***	1,3 (· · ·)

BEL	1990-1999	1999-2008
US gdp	-9,2 ***	-7,6 ***
L1	9,08	8,2 *** (· · ·)
L2	5,8	0,4
BEL gdp	0,4	9,7 * (·)
L1	11 *	-2,7 (·)
L2	-13 ***	-0,8 (· · ·)
US govt	-12,3 ***	-10,2 ***
L1	6,2	4,8 ** (··)
L2	9,7 **	3,3 ** (·)
BEL govt	0,1	5,7 *** (· · ·)
L1	30	0,2
L2	-27,3	-0,7
NER	0,5	1,04 *** (· · ·)
OECD gdp	2,5	-8,8 *** (· · ·)

NOTE: See note to Table 3.13

Table 24: Share of the reduction in the risk-sharing inefficiency between the US and the partner country explained by falls in the variance of shocks and improvements in the degree of insurance (CC preferences)

JAP	1978 break	1997 break	**CAN**	1997 break	
US gdp	-54,32	7,55	US gdp	0	
JAP gdp	-200,21	38,21	CAN gdp	0	
US govt	-17,86	9,72	US govt	-200,42	
JAP govt	44,59	15,68	CAN govt	-345,85	
NER	0	-8,85	NER	0	
OECD gdp	0	-9,25	OECD gdp	0	
Insurance	**-227,7%**	**53%**	**Insurance**	**-546,2%**	
Var(US gdp)	-11,62	2,21	Var(US gdp)	11,33	
Var(JAP gdp)	64,97	30,96	Var(CAN gdp)	311,6	
Var(US gov)	-11,42	1,09	Var(US gov)	34,29	
Var(JAP gov)	297,83	10,45	Var(CAN gov)	394,65	
Var(NER)	0	0	Var(NER)	0	
Var(OECD)	0	0	Var(OECD)	0	
C(gdps)	-11,97	2,24	C(gdps)	-105,61	
Shocks	**327,7%**	**46,9%**	**Shocks**	**646,2%**	
UK	1984 break	1999 break	**CH**	1983 break	1999 break
US gdp	-189,77	12,58	US gdp	-58,93	30,14
UK gdp	191,43	47,3	CH gdp	223,4	-0,39
US govt	78,68	-4,06	US govt	0	77,15
UK govt	79,88	-36,9	CH govt	19,04	1,96
NER	0	-8,4	NER	-0,62	-18,24
OECD gdp	0	0	OECD gdp	0	-1,02
Insurance	**160,2%**	**10,5%**	**Insurance**	**182,8%**	**89,6%**
Var(US gdp)	97,28	3,31	Var(US gdp)	61,89	5,35
Var(UK gdp)	29,34	84,34	Var(CH gdp)	-80,23	0
Var(US gov)	-111,33	1,84	Var(US gov)	-52,34	7
Var(UK gov)	2,32	0	Var(CH gov)	-15,32	-2,32
Var(NER)	-19,22	0	Var(NER)	0	0
Var(OECD)	0	0	Var(OECD)	0	0,37
C(gdps)	-31,62	0	C(gdps)	2,48	
Shocks	**-33,2%**	**89,4%**	**Shocks**	**-83,5%**	**10,4%**

NOTE: See note to Table 3.8

Table 25: Share of the reduction in the risk-sharing inefficiency between the US and the partner country explained by falls in the variance of shocks and improvements in the degree of insurance (CC preferences)

NDL	1996 break	**BEL**	1983 break
US gdp	-141,84	US gdp	-95,41
NDL gdp	0	BEL gdp	144,73
US govt	-190,91	US govt	99,48
NDL govt	0	BEL govt	-26,75
NER	591,52	NER	-64,5
OECD gdp	0	OECD gdp	-58,54
Insurance	**258,77**	**Insurance**	**-0,99**
Var(US gdp)	9,3	Var(US gdp)	13,88
Var(NLD gdp)	-111,48	Var(BEL gdp)	54,77
Var(US gov)	0	Var(US gov)	32,34
Var(NLD gov)	-204,62	Var(BEL gov)	0
Var(NER)	166,6	Var(NER)	0
Var(OECD)	0	Var(OECD)	0
Shocks	**-140,2**	**Shocks**	**100,99**
FIN	1978 break	**ITA**	1999 break
US gdp	-419,87	US gdp	35,42
FIN gdp	463,39	ITA gdp	-6,6
US govt	170,9	US govt	0
FIN govt	116,89	ITA govt	47,77
NER	0	NER	0
OECD gdp	0	OECD gdp	0
Insurance	**331,31**	**Insurance**	**76,59**
Var(US gdp)	-133,37	Var(US gdp)	-1,3
Var(FIN gdp)	182,24	Var(ITA gdp)	0
Var(US gov)	-461,97	Var(US gov)	1,98
Var(FIN gov)	91,13	Var(ITA gov)	29,46
Var(NER)	0	Var(NER)	6,12
Var(OECD)	93,47	Var(OECD)	0
Shocks	**-228,5**	**Shocks**	**36,26**

NOTE: See note to Table 3.8