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Abstract

In spite of two decades of financial globalization, consumption-based indicators do not seem to signal more international risk sharing. We argue that consumption risk sharing among industrialised countries *has* actually increased – in particular since the 1990s – but that standard consumption-based measures of risk sharing – such as the volatility of consumption conditional on output or international consumption correlations – have been unable to detect this increase. The reason is that consumption has also been affected by the concurrent decline in the volatility of output growth in most industrialised countries since the 1980s. As a first important driver of this decline we identify a more gradual response of output to permanent idiosyncratic shocks. Since consumption reacts mainly to permanent shocks, it appears more volatile in relation to current changes in output. This effect seems to have offset the tendency of financial globalization to lower the volatility of consumption conditional on output. Secondly, because the variability of permanent global shocks has also fallen, international consumption correlations have also generally not increased as financial markets have become more integrated.

KEYWORDS: *Consumption Risk Sharing, International Business Cycles, Great Moderation, Financial Integration and Capital Flows, Home Bias*

JEL CLASSIFICATION: C23, E21, F36

1 Introduction

Since the beginning of the 1980s, international financial markets have become increasingly integrated. It would seem that this should unambiguously lead to better international consumption risk sharing. But conventional measures of international consumption risk sharing seem to be unaffected by more than two decades of financial globalization and continue to display low levels of financial market integration. In this paper, we document that consumption risk sharing has actually improved during the globalization period but that the impact of financial globalization on consumption-based measures of risk sharing may have been blurred by the fact that the underlying risks have changed as well: international business cycles have become much less volatile.

The literature on consumption risk sharing falls into two broad strands. The first strand emphasizes that in complete financial markets, marginal utility growth should be equated across countries so that consumption growth rates should be highly correlated. This correlation-based approach has encountered the now famous consumption correlation or quantity puzzle: in the data the correlation in consumption is not only low (consistent with low degrees of international risk sharing) but also generally lower than the correlation in the underlying risks, i.e. output growth rates.(Backus, Kehoe and Kydland (1992)).

A second strand of the empirical consumption risk sharing literature has focused on regression-based measures. These studies (notably Asdrubali, Sørensen and Yosha (1996), Sørensen and Yosha (1998) and Crucini (1999)) emphasize an alternative prediction of the complete markets model: fluctuations in relative (i.e. idiosyncratic) marginal utility growth should be independent of idiosyncratic risk (as measured by relative output growth

rates). Therefore, the coefficient of a (panel-) regression of relative consumption growth on relative output growth should be zero.[1] Similar risk-sharing regressions were first proposed by Mace (1991), Cochrane (1991) and Townsend (1991) as tests of the null of market completeness, but as argued very convincingly by Asdrubali, Sørensen and Yosha (1996) they are more generally useful: since the estimated coefficient is typically between zero and one it can directly be interpreted as the fraction of risk that remains unshared: if the coefficient is unity, no risk is shared, if it is zero, all risk is shared.

Both risk sharing regressions as well as consumption correlations document a lack of risk sharing in international relative to regional data, but they do not seem to change very much *over time*. None of the papers cited above detects a major increase in risk sharing in international data in what we refer to as the globalization period, i.e. after 1980, even though international capital flows as well as cross-holdings of equity and foreign direct investment have seen spectacular growth (see Lane and Milesi-Ferretti (2001, 2004)). Recently, Moser et al. (2003) have investigated risk sharing regressions based on European data and concluded that consumption risk sharing has not increased since capital markets have been liberalized nor since the move to a common currency. Based on a broad sample of 43 highly financially integrated economies, Bai and Zhang (2005) do not find a significant increase in consumption risk sharing either. Heathcote and Perri (2004) even document a marked *decrease* in the correlation between U.S. consumption and consumption in other industrialised economies. Labhard and Sawicki (2006), using UK regional and international data, equally find

[1] We use the term 'relative' interchangeably to denote the idiosyncratic component of a variable constructed.as the logarithmic difference between a country-realization of that variable and its cross–sectional mean.

a slight decrease in risk sharing based on a factor-analytical approach.

All these findings could be taken as evidence that financial integration has not helped to facilitate international consumption risk sharing. In this paper we argue that this interpretation is incorrect. In fact, consumption risk sharing has increased but consumption-based measures of risk sharing have mostly been unable to detect the impact of financial globalization because of the concurrent decline in the volatility of output growth in most industrialised countries.[2]

Specifically, we argue that risk sharing regressions have not picked up the effect of financial globalization because the volatility of country-specific output fluctuations has dropped by more in the short-run than in the long-run, reflecting what appears to be a more gradual response of output to country-specific shocks. Since, as we show, consumption reacts primarily to permanent changes in relative output, the volatility of relative consumption conditional on *current* relative output growth has not decreased. While most of our discussion centers on risk sharing regressions, we also show that consumption growth correlations also do not seem to have increased because the volatility of the global component of permanent output shocks has decreased even more sharply than has the volatility of their country-specific components.

We discuss the impact of these changes in the volatility of business cycles in the framework of a permanent income model in which all transitory fluctuations in consumption can be smoothed but in which there is also some partial, though incomplete, insurance against permanent idiosyncratic

[2]Since the first version of our paper (Artis and Hoffmann (2004)), a small number of studies has started to find similar indications that risk sharing might have increased with financial integration (see Bekaert et al. (2006) for emerging market economies following liberalizations and Sørensen et al (2007) for the link between home bias and risk sharing), but to date these remain exceptions (see also the excellent survey in Kose et al. (2007)).

shocks. From a theoretical perspective, the distinction between permanent and transitory fluctuations in output and income is central, because it is plausibly much harder for households, regions and entire countries to insure against permanent rather than against transitory fluctuations: Insurance against permanent shocks requires *ex-ante* diversification which is generally only possible through state-contingent assets such as equity, whereas transitory variation in income can also be smoothed *ex-post* through borrowing and lending.[3]

In our theoretical framework we show that the coefficient of the regression of relative consumption growth on current relative output growth will not only depend on the degree of partial insurance – which is the parameter through which we capture the degree of financial globalization – but also on the dynamic pattern of adjustment to shocks. To see the intuition for this, consider two countries both of which have the same degree of partial insurance and both of which experience a permanent country-specific shock of the same size. However, we assume that in the first country, the adjustment to output to the new permanent level is immediate, whereas in the second, output adjusts only gradually to its new long-run level. In the first case, the current change in output corresponds to the permanent change in output. The regression of consumption on output growth therefore correctly reveals the degree of insurance: it tells us how much of the volatility in permanent output systematically spills over into consumption. In the second case, however, consumption will actually react more strongly than current output because it immediately jumps to the level implied by the permanent change

[3]Markets for state contingent assets are more likely to be subject to frictions and market incompleteness than are markets for non-state contingent assets such as bonds or loans. State-contingent assets will only exist to the extent that the state of the world, on which they are contingent, is not too costly to verify. As pointed out in e.g. Kocherlakota (1996), problems of moral hazard or enforceability are particularly likely to arise in such markets which may render them endogenously incomplete.

in output, whereas current output follows suit only over time. Hence, even though both countries are equally well insured, the risk sharing regression will tend to underestimate the actual amount of risk sharing for the second country by delivering a coefficient that appears too high.

In our empirical analysis we show that this mechanism – an instance of Deaton's (1992) paradox – can indeed explain why the coefficient of conventional risk sharing regressions has not fallen as financial globalization has advanced. Our results are based on data from 22 OECD countries spanning the period 1960-2000. For the sake of comparison, we also obtain results for U.S. state level data for the period 1960-90. As is suggested by our theoretical framework, we account separately for permanent (trend growth) and transitory (purely cyclical) variation in country-specific business cycles. We first confirm that transitory shocks have indeed little or no impact on consumption, whereas insurance against permanent shocks is much less complete, very much as the permanent income model with partial insurance would suggest. Most importantly, however, we detect a considerable increase in the degree of international risk sharing during the globalization period – i.e. during the 1980s, and even more so during the 1990s. This increase is associated mainly with a better diversification of permanent risk. The reason why this effect, which is most likely the result of financial globalization[4], has not been picked up by conventional risk sharing regressions is that there is a marked moderation in the volatility of output vis-à-vis the pre-globalization era. This moderation is associated with a more gradual response of the country-specific component of output to permanent shocks, which gives rise to the Deaton-effect discussed in the example and which

[4] This increase in risk sharing is indeed closely linked with the growth in international asset positions (see our companion paper Artis and Hoffmann (2007) and also Sørensen, Wu, Yosha and Zhu (2007)).

induces the risk sharing regression to underestimate the true, considerably increased, extent of international risk sharing.

Our results tie in with a recent theoretical and empirical literature that emphasizes that financial globalization may in itself have important effects on international business cycles. In particular, they are in line with the observation that international business cycles have become more synchronized over the last two decades and that business cycle volatility has decreased globally – a phenomoneon often referred to as the great moderation (see McConnel and Perez Quiros (2000), Kose et al. (2003, 2005), Bordo and Helbling (2004) and Stock and Watson (2005)). Our paper is also related to the work of Imbs (2006) who provides cross sectional evidence that financial integration seems to lead to more business cycle symmetry and that it may increase risk sharing. But Imbs also acknowledges that the consumption correlation puzzle does not seem to vanish and that changes in consumption based measures of international risk sharing over time are notoriously hard to detect. Our results suggest a solution to this puzzle.

Our framework and empirical findings also bear some interesting similarities to a recent micro-labour literature that emphasizes the distinction between permanent and transitory shocks for understanding household level consumption inequality. Notably, Heathcote, Storesletten and Violante (2007) obtain closed-form solutions to an incomplete markets model that can serve as an analytical framework in which any degree of partial insurance can be allowed for. Based on a similar line of argument as we have used here, Blundell, Pistaferri and Preston (2006) show that household consumption inequality in the U.S. has not increased in line with income inequality mainly because the relative contribution of transitory shocks to income inequality has increased and because it is easier for households to obtain insurance

against the transitory part of their idiosyncratic risk.

The remainder of this paper is structured as follows: in section two we discuss both the regression- and correlation-based approaches to the measurement of consumption risk sharing. Section three lays out the theoretical framework of the partial insurance model based on which we then propose a decomposition of the conventional risk sharing coefficient into one component that captures the dynamic adjustment of output to shocks and a (business cycle-) adjusted measure of risk sharing. In section four we present our data and the details of the empirical implementation. Section five offers a discussion of our main results. While we find it useful to discuss most results in this paper based on a regression-based approach, this section also applies our framework to the analysis of consumption correlations. Section six concludes.

2 Consumption-based measures of risk sharing

The consumption-based measures of risk sharing that we consider in this paper and that are most used in the literature are motivated by a benchmark model with complete financial markets and frictionless trade in goods. In such a model, marginal utility growth in country or region k equals the growth in the shadow price of consumption and is therefore equalized across countries:

$$\frac{u'_k(C^k_{t+1})}{u'_k(C^k_t)} = \frac{\mu_{t+1}}{\mu_t} \tag{1}$$

where $u'(.)$ is the period utility function and C^k_t measures consumption in country k. The shadow price of consumption is μ_t. There are two related readings of this fundamental equation that have both found their reflec-

tions in the empirical literature. The first is that marginal utility growth should be perfectly correlated across countries. One branch of the literature therefore looks at consumption correlations. This line of research has encountered the now famous consumption correlation puzzle (Backus, Kehoe, Kydland (1992)), the stylized fact that international consumption correlations are lower than the corresponding output correlations. Stockman and Tesar (1995) have argued that consumption is likely to be driven by preference shocks and subject to considerable measurement error so that low empirical consumption correlations could in principle arise even in complete markets. This can make it difficult to interpret consumption correlations as *measures* of risk sharing.

Our main focus in this paper will therefore be on an alternative reading of equation (1) that has equally made a profound impact on the risk sharing literature: since growth in the shadow price is common to all countries, the difference between marginal utility growth in two countries, while not necessarily zero, should be statistically independent of the country-specific risk-variables, notably relative endowments.

In order to obtain an estimable equation, specific assumptions on the form of the utility function are typically made. Under logarithmic utility, the optimality condition can be written

$$\mathbf{E}\left[\Delta c_t^k - \Delta c_t^* | \mathbf{X}_t^k\right] = 0$$

where $\mathbf{X}_t^k$ is a vector of time-varying country characteristics that capture idiosyncratic risk, such as relative output growth; lower case letters denote logarithms, so that Δc_t^k is the growth rate of consumption in country k. The asterisk denotes foreign (world average) variables.

Under full insurance, the regression

$$\Delta c_t^k - \Delta c_t^* = \mathbf{b}'\mathbf{X}_t^k + \varepsilon_t$$

should yield a coefficient of zero.[5]

Unlike consumption correlations, the regression-based formulation of the test for full risk sharing allows consumption to be affected by a host of unobserved factors, such as measurement error and preference shocks. Provided these factors are uncorrelated with the idiosyncratic risk characteristics $\mathbf{X}_t^k$, they should just be captured by the regression residual.

Clearly, $\mathbf{X}_t^k$ could itself be correlated with relative consumption growth for reasons that are unrelated to financial market incompleteness. For example, non-separabilities in utility between consumption and leisure will render income and consumption endogenous and correlated. However, one of the central messages of the international real business cycle literature (notably Backus, Kehoe and Kydland (1992)) is that the substitution effect between consumption and leisure can at best account for a very limited degree of comovement in relative consumption and output at the aggregate level. Another instance in which one might expect a correlation between $\mathbf{X}_t^k$ and relative consumption is if $\mathbf{X}_t^k$ is related to relative movements in consumption price levels. The first order condition (1) assumes that all countries consume the same bundle of goods and that these goods are freely traded so that consumption price levels equalize. If this is not the case, then complete risk sharing implies a high correlation between relative consumption and the real exchange rate (Backus and Smith (1993)). Here, we disregard the role of real exchange rate fluctuations[6] and follow the bulk of the literature

[5] Mace (1991) and Cochrane (1991) were the first authors to investigate regression of this type in household-level data.

[6] Hoffmann (2008) reports that deviations from purchasing power parity can indeed

that has treated $\mathbf{X}_t^k$, notably relative output, as exogenous in the empirical analysis. Specifically, most researchers, including Asdrubali, Sørensen and Yosha (1996) and Crucini (1999) have formulated regressions of the form

$$\Delta c_t^k - \Delta c_t^* = b\left[\Delta y_t^k - \Delta y^*\right] + \varepsilon_t \tag{2}$$

where y^k is the logarithm of output in country k and the asterisk, again, denotes the world average. We call this equation the basic risk sharing regression. Clearly, in models with complete markets (and against the backdrop of the qualifications given above) the coefficient estimate of b should be zero or close to zero. Real world financial markets are likely to be incomplete and panel estimates of b are typically between zero and one. This has led researchers, notably, Asdrubali, Sørensen and Yosha (1996), to interpret b as a *measure* of risk sharing that indicates the fraction of idiosyncratic risk that remains unshared. Based on US state level data, Asdrubali, Sørensen and Yosha find that b is around 0.25, which suggests that roughly a quarter of idiosyncratic output fluctuations remain uninsured. Based on data from OECD countries, Sørensen and Yosha (1998) find more than 70 percent of idiosyncratic fluctuations appear to remain uninsured. Hence, according to this measure, there is a lack of international consumption risk sharing when risk sharing within countries is taken as the benchmark.

What is puzzling, however, is that in spite of the recent wave of financial globalization, regressions of the form (2) do not seem to indicate an increase in international risk sharing – the coefficient b does not seem to decline. We

account for a substantial fraction of the comovement between relative consumption and relative output. But he also shows that the conclusions as to how much risk is shared through financial markets remain unaffected by controlling for real exchange rate movements. It therefore seems an acceptable simplification to abstract from the role of exchange rate movements in answering our basic question here: why has risk sharing not increased as financial portfolios have become more globalized?

argue next that international business cycles have changed in a way that has offset the tendency of b to decline with better financial integration.

3 A permanent income model with partial insurance

We use a simple theoretical framework inspired by Crucini (1999) in which we assume that a country can obtain insurance against permanent (idiosyncratic) fluctuations in its income by buying into a world mutual fund that pays world average output as a dividend. Therefore, income growth is a weighted sum of domestic and world average output growth:

$$\Delta inc_t^k = (1-\omega)\Delta y_t^k + \omega \Delta y_t^*$$

Here, ω measures the fraction of country k's wealth held in the mutual fund. This measure of diversification is our index of financial globalization. We further assume that, once income is observed, the country can fully smooth the effect of any transitory fluctuations in consumption through borrowing and lending so that

$$\Delta c_t^k = \Delta inc_t^{kP} = (1-\omega)\Delta y_t^{kP} + \omega \Delta y_t^{*P} \tag{3}$$

where the superscript 'P' denotes the permanent component.

Acknowledging that for the world as whole it must be that $\Delta c^* = \Delta inc^{*P} = \Delta y^{*P}$, we can rewrite equation (3) as

$$\Delta c_t^k - \Delta c_t^* = (1-\omega)\left[\Delta y_t^{kP} - \Delta y_t^{*P}\right] \tag{4}$$

If $\omega = 0$, then there is no international diversification and the country ob-

tains no insurance against permanent idiosyncratic output shocks, whereas if $\omega = 1$, insurance against such shocks will be complete. In general, we should expect ω to be between 0 and 1, implying what we call imperfect risk sharing or partial consumption insurance.

This model provides guidance in that it suggests the importance of distinguishing between permanent and transitory fluctuations in the empirical measurement of risk sharing. First, equation (4) would suggest a regression of relative output growth rates on relative growth rates in permanent output:

$$\Delta c_t^k - \Delta c_t^* = b_P \left[\Delta y_t^P - \Delta y_t^{*P}\right] + \xi_t \tag{5}$$

and under the null of our model, $b_P = (1 - \omega)$. Secondly, the coefficient b_T in the regression

$$\Delta c_t^k - \Delta c_t^* = b_T \left[\Delta y_t^T - \Delta y_t^{*T}\right] + v_t \tag{6}$$

where $\Delta y^T = \Delta y - \Delta y^P$, should equal zero - transitory variation in relative outputs should not have an impact on relative consumption. For empirical purposes, requiring that b_T is zero may be too strict a condition to impose. After all, our theoretical assumption that transitory fluctuations can actually be smoothed away completely whereas insurance against permanent shocks is generally incomplete is just a metaphor for saying that existing financial markets make it harder for countries to insure against permanent shocks than against transitory fluctuations. Our empirical analysis specifically allows for the possibility that b_T is different from zero, for example because a fraction of consumers faces borrowing constraints. But our expectation would be that $b_p = (1 - \omega) > b_T$ – a conjecture we find confirmed in

the data.

We consider b_P and b_T to be the 'structural' risk sharing coefficients. As we show next, the coefficient of the basic risk sharing regression (2), b, can be interpreted as a reduced form that is not only a function of these structural coefficients b_P and b_T, but also of what we call the structure of business cycles: the respective contributions of permanent and transitory fluctuations to the volatility of output and the shape of the dynamic response to permanent shocks.

3.1 Link to the basic risk sharing regression

In the remainder of the paper we economize on notation and let the hat denote relative growth rates of a variable. In addition, whenever possible without ambiguity, we drop the time subscript t so that $\widehat{c} = \Delta c - \Delta c^*$ and $\widehat{y} = \Delta y - \Delta y^*$. Then the regression coefficient b of the baseline risk sharing regression is

$$b = \frac{cov(\widehat{c}, \widehat{y})}{var(\widehat{y})} = \frac{cov(\widehat{c}, \widehat{y}^P) + cov(\widehat{c}, \widehat{y}^T)}{var(\widehat{y})}$$

whereas the two structural risk sharing coefficients from (5) and (6) are given by

$$b_P = \frac{cov(\widehat{c}, \widehat{y}^P)}{var(\widehat{y}^P)} \text{ and } b_T = \frac{cov(\widehat{c}, \widehat{y}^T)}{var(\widehat{y}^T)}$$

We can then write

$$b = b_P \frac{var(\widehat{y}^P)}{var(\widehat{y})} + b_T \frac{var(\widehat{y}^T)}{var(\widehat{y})} \tag{7}$$

This equation allows us to infer that a drop in the structural risk sharing coefficients can be offset by changes in the ratios $var(\widehat{y}^P)/var(\widehat{y})$ and $var(\widehat{y}^T)/var(\widehat{y})$. In the data, we find that b_T is generally close to zero, quite

in line with our model of partial insurance. In our discussion of the impact of changes in business cycle volatility on b, we therefore find it useful to impose the restrictions from our model, specifically $b_T = 0$ and $b_P = (1 - \omega)$. This simplifies (7) to

$$b = (1 - \omega)\phi \tag{8}$$

where $\phi = var(\widehat{y}^P)/var(\widehat{y})$. We call ϕ the long-term variance ratio.

The decomposition (8) allows us to identify the sources of variation in b over time. Based on international data, we document that ω has actually increased — risk sharing has actually improved a lot – but an increase in ϕ has offset the impact of better international diversification on b.

We interpret ϕ as an indicator of the pattern of dynamic adjustment following a country-specific shock to the permanent component of (relative) output. This is most easily demonstrated in the context of an AR(1) process for country-specific output growth. Let $\widehat{y}_t = \rho\widehat{y_{t-1}} + \varepsilon_t$ and $var(\varepsilon_t) = \sigma^2$. Then the permanent component of a shock ε_t is $\varepsilon_t^P = (1 - \rho)^{-1}\varepsilon_t$ and $var(\varepsilon_t^P)/var(\varepsilon_t) = 1/(1 - \rho)^2$. An increase in ρ will increase the long-term variance ratio ϕ and is associated with a more sluggish and gradual adjustment of output to its new long-run level.

We further illustrate the role of ϕ in figures 1 and 2. In both figures there is a shock to permanent output $\widehat{y}^P$ of the same size. In both cases, relative consumption will therefore immediately adjust to its new permanent level so that $\widehat{c} = (1 - \omega)\widehat{y}^P$. However, in figure 1, the adjustment of output to the permanent shock is immediate, so that $\widehat{y} = \widehat{y}^P$, whereas in figure 2 output adjusts only gradually to its new long-run level, so that $\widehat{y} < \widehat{y}^P$. In the first case, we have $\phi = 1$ and the regression of $\widehat{c}$ on $\widehat{y}$ will correctly identify the degree of risk sharing, $(1 - \omega)$. In the second figure

though, since $\widehat{y} < \widehat{y}^P$, consumption will actually appear excessively volatile relative to current output because it fully reacts to the permanent shock on impact, whereas current output follows suit only over time. Hence we have $\widehat{c} = (1-\omega)\widehat{y}^P = (1-\omega)\phi\widehat{y}$ with $\phi > 1$. This instance of Deaton's (1992) paradox causes the basic risk sharing regression to underestimate ω. But note that this un-smoothing of consumption is an optimal response of the consumer to the permanent shock, not a failure to share consumption risk.[7]

We find that the increase in ϕ is directly linked to the general decrease in business cycle volatility over the 1980s and 1990s. In our data, both $var(\widehat{y}^P)$ and $var(\widehat{y})$ have decreased, but the variability at the business cyclce frequency, $var(\widehat{y})$, has actually fallen more than has $var(\widehat{y}^P)$. While it is not the aim of this paper to explore the economic causes of changes in business cycle volatility, our results are in in line with many studies in the area in indicating that shocks to output have not only become less volatile but that the response of output to these shocks seems to have become more gradual. Possibly this is the reflection of better macroeconomic management and a generally increased resilience of advanced economies to idiosyncratic shocks.

4 Econometric implementation

4.1 Constructing permanent components

Estimating our set of risk sharing regressions (5) and (6) involves the identification of a permanent component of domestic and foreign output growth. We construct this as the change in the annuity-value of domestic or foreign output according to

[7] Since it is not empirically relevant in our data set and in order to save on space, the figure does not provide the third case in which $\phi < 1$ so that current output changes exceed the long-run changes. In analogy to case 2, this would lead us to overestimate risk sharing.

$$Y_t^P = (1-R)\sum_{k=0}^{\infty} R^k \mathbf{E}\left(Y_{t+k}\right) \tag{9}$$

where $R = (1+r)^{-1}$ and r is the world real interest rate. In the appendix, we derive the following approximation for the logarithm of Y_t^P:

$$y_t^P = \log\left[Y_t^P\right] \approx y_t + \sum_{k=1}^{\infty} R^k \mathbf{E}\left(\Delta y_{t+k}\right) = (1-R)\sum_{k=1}^{\infty} R^k \mathbf{E}\left(y_{t+k}\right) \tag{10}$$

Once we specify processes for Δy_t and $\Delta y^*$to proxy for the expectations involved in (10) we can then obtain $\widehat{y}_t^P = \Delta y_t^P - \Delta y_t^{*P}$ and $\widehat{y}_t^T$.

We follow Crucini (1999) and consider several different specifications for the stochastic processes driving Δy and Δy^*: in the first, home and foreign output follow separate AR(1) processes whereas in the second, home and foreign output and consumption follow a VAR. We then use the estimated processes to compute the expectations involved in the construction of Δy^P and Δy^{*P} in (10).

We describe the construction of these permanent values in the appendix. We also performed a battery of robustness checks on our identification of permanent components. Specifically, we experimented with values of r between 0 and 0.1 and we also let r vary across countries to reflect the effects of limited financial integration. Our findings turned out to be extremely robust. In the remainder of the paper, we present results that are based on the same value of $r = 0.02$ for all countries and regions.

The permanent-transitory decompositions in the spirit of Beveridge and Nelson (1981) as we have just described them give us a permanent component Y^P that is virtually a random walk, irrespective of what the driving

process for Y_t is.[8] This feature of Beveridge-Nelson-type decompositions is central for our analysis: in theory, what should matter for consumption dynamics are *unpredictable* movements in the trend. This is the reason why in a permanent-income setting, permanent and transitory fluctuations in income are qualitatively different with respect to their degree of insurability: ultimately, permanent fluctuations in output *cannot* be smoothed through borrowing and lending and therefore have to be insured through more sophisticated assets such as equity.[9] We therefore restrict our analysis here to trend-cycle decompositions that are directly derived from and consistent with the underlying economic theory.[10]

4.2 Data and estimation

We use annual data from the Penn World Table, release 6.1 (PWT 6.1.) by Heston, Summers and Aten (2002). The sample ranges from 1960 to 2000. All data are in constant (1996) international prices. The countries included in our estimation are:

1. Canada, 2. the United States, 3. Japan, 4. Austria, 5. Belgium, 6. Denmark , 7. Finland, 8. France, 9. Germany (West), 10. Greece, 11. Iceland, 12. Ireland, 13. Italy, 14. Netherlands, 15. Norway, 16. Portugal, 17. Spain, 18. Sweden, 19. Switzerland, 20. United Kingdom, 21. Australia, 22. New Zealand. Most of these countries are OECD countries and we will refer to them under this label.

[8] Note that in the case $r = 0$, that we consider in the range of possible values for the discount factor, we obtain exactly the Beveridge-Nelson permanent component.

[9] See e.g. Baxter and Crucini (1995) who show that the dynamic properties of bonds-only and complete markets economies are different only for very persistent shock processes.

[10] Specifically, we do not apply a battery of purely statistical filtering methods. For example, the permanent component obtained from an HP-filter is smooth for standard parameter values and this smoothness implies a high degree of predictability. Therefore changes in the HP permanent component will not generally correspond closely to changes in the theoretically relevant permanent component which is the annuity value of output.

Our main interest here is in using our decomposition of the basic risk sharing coefficient to understand how the measurement of risk sharing is affected over time. However, one might also want to ask to what extent differences in the structure of international and regional business cycles (as captured by differences in ϕ) can account for why there is an apparent lack of risk sharing among countries relative to regions within a country. By way of comparison, we therefore also apply our analysis to the U.S. state level data set used in the study by Asdrubali, Sørensen and Yosha (1996)[11]. This data set ranges from 1960-90.

For the international data set we report results for three subperiods: the first covers the period 1960-1980, the second covers 1980-2000. Finally, we also look at the 1990s alone. There are a number of reasons for choosing these particular sub-periods. First, the 1980s saw the major liberalization of capital markets and we therefore refer to the post-1980 part of our sample as the globalization period. Secondly, reporting separate results for the 1990s alone will provide insights into the effects on risk sharing of the dramatic increase in net international asset positions that started to take place from the early 1990s onwards (compare e.g. the data in Lane and Milesi-Ferretti (2001)). Third, the sample split after 1980 is also justified on the grounds that the properties of international business cycles have started to change dramatically in the 1980s, a phenomenon often referred to as the great moderation.[12]

[11] The data base is available at Oved Yosha's web page `http://econ.tau.ac.il/research/riskshare/channels/channels.htm`

We do not describe the data here since it is described in detail in the orginal article as well as in Becker and Hoffmann (2006) and a previous version of this paper (Artis and Hoffmann (2004)).

[12] We also obtained results for the period 1960-90, because this is the period covered in the influential studes by Sørensen and Yosha (1998) and Crucini (1999) and in the U.S. data set by Asdrubali, Sørensen and Yosha (1996). We do not report these results though, since they are qualitatively very similar to those obtained for the 1960-80 period.

We estimated all risk sharing regressions with a panel two-stage least squares procedure. First, we removed country-specific fixed effects. Then we estimate the panel by OLS. To control for heteroskedasticity, all variables are then weighted by the country-specific variance of the first stage residuals and the model is re-estimated.

5 Empirical results

The first line in table 1 provides the results of basic risk sharing regressions for both U.S. and international data. Roughly three quarters of idiosyncratic output variability remains uninsured in country-level data, in the later period (1980-2000), more than 80 percent. Only 15 percent of idiosyncratic variability spills over into consumption according to the results obtained from U.S. state level data.

Our basic risk sharing regressions clearly reflect the general pattern in the literature: first, according to b, there is a lot more risk sharing in regional data than there is in country-level data, but even at the regional or state level, risk sharing is not complete. Secondly, in international data the estimate of b does not seem to decline. For the period 1980-2000 it is actually *higher* than that obtained for the earlier period. The point estimate for the 1990s is marginally lower, but still higher than for the 1960-80 period and not significantly different from that for the entire 1980-2000 period. From these estimates it is clear that the effects of financial globalization do not seem to work through to the risk sharing regression.

These results constitute our point of reference. We will refer back to table 1 as the results of the 'baseline specification', or equivalently, of the 'basic risk sharing regression'. We now turn to dissecting these results.

5.1 Insurance of permanent and transitory shocks

In table 2 we present the results from the risk sharing regressions on permanent and transitory output variation. In U.S. data we find that only 5-15 percent of permanent variability remain uninsured. A similar value obtains for transitory fluctuations. Hence, in U.S. data we cannot find evidence that there is a qualitative difference between permanent and transitory shocks to output in so far as their degree of insurability is concerned.[13]

The picture changes quite substantially once we turn to the regression with international data. While OECD countries also seem to be able to smooth transitory variation almost completely, at least in the later subperiods, the coefficient on permanent output variation tells us that in 1960-80 only little more than 40 percent of permanent idiosyncratic output variability were insured at the international level – b_P is around 0.57. But this coefficient has come down quite markedly over time. In the period 1980-2000 it is 0.43 (0.42 for the AR) already considerably lower than for the 1960-80 period even though the difference is only marginally significant. In the 1990s there is a further dramatic drop that is now also clearly significant vis-à-vis the 1960-80 period.

While there is still a lot less insurance at the international than at the regional level, the results here are indicative of a substantial increase in international consumption risk sharing over the globalization period. The baseline regression does not seem to pick this up.

[13] This result is in line with earlier findings by Asdrubali, Sørensen and Yosha (1996) who document that idiosyncratic persistence does not seem to have a big effect on the overall extent of insurance in U.S. data but that regions with more persistent idiosyncratic fluctuations rather tend to insure *ex-ante*.

5.2 Effects of business cycle structure

Table 3 sheds light on why the coefficient b seems unaffected by financial globalization.

As is apparent from the first colums, the volatility of the country-specific component of business cycles, $var(\widehat{y})$, has fallen considerably over the three sub-periods.[14] Since we can also write $var(\widehat{y}) = var(\Delta y) - 2cov(\Delta y, \Delta y^*) + var(\Delta y^*)$, this is in line with the observation made by many authors that international business cycles have both become less volatile and more synchronized.[15] Our estimates for $var(\widehat{y}^P)$, reported in the second column, show that this decrease in idiosyncratic volatility is also associated with a marked drop in the volatility of the permanent idiosyncratic component in output growth. But though $var(\widehat{y}^P)$ has also decreased quite substantially, $var(\widehat{y})$ has fallen by more, as is reflected by the marked growth of the long-term variance ratio ϕ reported in the third column.

In column 4 we report what we call the business-cycle adjusted risk sharing coefficient

$$b_{adj} = b/\phi \tag{11}$$

which is the baseline risk sharing coefficient from table 1 normalized by the long-run variance ratio. It is readily apparent that

$$b_{adj} = b_P + b_T \frac{var(\widehat{y}^T)}{var(\widehat{y}^P)} = (1-\omega) + b_T \frac{var(\widehat{y}^T)}{var(\widehat{y}^P)}$$

We think about b_{adj} as a compound measure of risk sharing that explicitly allows for the possibility that both consumption insurance and consumption

[14] We calculate the variances $var(\widehat{y})$, $var(\widehat{y}^P)$ and $var(\widehat{y}^T)$ as the cross-country average of the variances of the respective variable, so that e.g. $var(\widehat{y}) = \frac{1}{K}\sum_{k=1}^{K} var(\widehat{y}_t^{\,k})$ and $var(\widehat{y}_t^k)$ is the time-series variance of $\widehat{y}$ for country k.

[15] See Perez, Osborn and Artis (2006) for a recent contribution in this mould and the references cited there for a more comprehensive survey of the literature.

smoothing could be imperfect, so that $b_P = (1-\omega)$ and b_T are both positive. Since $b_T \frac{var(\widehat{y}^T)}{var(\widehat{y}^P)} \geq 0$, the difference $b_{adj} - b_P$ will be positive and increasing in the importance of transitory relative to permanent variation in country-specific output. In the data, we have found that b_T is reasonably small and generally insignificant. By imposing $b_T = 0$ we can therefore infer a lower bound $\overline{\omega}$ on the globalization parameter ω by associating $b_{adj} = (1 - \overline{\omega})$.

Our estimates of the business-cycle adjusted risk sharing coefficient convey a clear message: given the rise in the long-term variance ratio, risk sharing must have increased; b_{adj} drops from 0.7 in 1960-80 to 0.5 for the globalization period as a whole and continues its decline in the 1990s with a drop to below 0.4 and the decline from each period to the next appears significant. If we impose $b_T = 0$, the values implied for $\overline{\omega}$ suggest that the fraction of permanent idiosyncratic risk insured at the international level must have increased from 30% to at least 60% in the 1990s.

By way of comparison, the last row of table 3 reports results for the United States. As is apparent, US federal states do not have a systematically lower long-term variance ratio that could help explain why the estimates of b are so much lower in regional than in international data. While this is not the central part of our analysis, we emphasize this point, because it implies that the conclusions that a vast literature has drawn from comparing estimates of b across space remain unaffected: When compared to the U.S., there remains a clear lack of risk sharing at the international level. But our results here suggest that the gap between how much risk is shared within a country and between countries has, though not closed, so clearly narrowed substantially.

Before we conclude, we now turn to an examination of the role of changes in the synchronization and volatility of international business cycles for con-

sumption correlations.

5.3 Implications for consumption correlations

The first part of table 4 presents cross-country averages of international consumption and output growth correlations for both the U.S. and for the OECD countries – for the latter again for the three subperiods. As is apparent, consumption correlations for the U.S. are not that much higher than international correlations. Both international *and* regional correlations indicate a significant lack of risk sharing in that they are much lower than unity. In addition, in both data sets output correlations are actually higher than consumption correlations. This is the quantity puzzle that, as we have argued, limits the interpretability of consumption correlations as a measure of risk sharing. Still, one should expect that financial integration would have led to an increase in international consumption correlations. In table 4 however, there is clearly no evidence of such an increase. Very much as in the case of the risk sharing regression, changes in world business cycles may have offset the effect of financial integration on international consumption correlations.

Our partial insurance framework implies a simple factor structure for domestic and international consumption that also allows us to explore the impact of changes in business cylce volatility on international consumption correlations:

$$\begin{bmatrix} \Delta c_t^* \\ \Delta c_t \end{bmatrix} = \begin{bmatrix} 1 & 0 \\ \omega & 1-\omega \end{bmatrix} \begin{bmatrix} \Delta y_t^{*P} \\ \Delta y_t^P \end{bmatrix} = \begin{bmatrix} 1 & 0 \\ 1 & 1-\omega \end{bmatrix} \begin{bmatrix} \Delta y_t^{*P} \\ \Delta y_t^P - \Delta y_t^{*P} \end{bmatrix} \tag{12}$$

where Δy_t^{*P} can be interpreted as the 'common factor' in domestic and

foreign consumption growth. To the extent that the global and country-specific factors are orthogonal, (12) implies that changes in consumption correlations can come about either due to changes in the variance of common relative to idiosyncratic shocks to Δy_t^P or due to changes in in ω.

Now assume that between the pre-globalization and the globalization periods, the variance of the the common relative to the idiosyncratic factors has changed so that

$$\left[\frac{var(\Delta y_t^{*P})}{var(\widehat{y}_t^P)}\right]_{t>t_0} = (1+\delta)\left[\frac{var(\Delta y_t^{*P})}{var(\widehat{y}_t^P)}\right]_{t<t_0}$$

for some point in time t_0. Under the assumption that ω has remained the same between periods, Forbes and Rigobon (2001) have shown, that the correlation coefficient implied by for the period after the change can be written as

$$\rho_{imp} = \rho_{_{t<t_0}}\sqrt{\frac{1+\delta}{1+\delta\rho_{_{t<t_0}}^2}} \tag{13}$$

where $\rho_{_{t<t_0}}$ is the correlation coefficient from the period before the change in $var(\Delta y_t^{*P})/var(\widehat{y}_t^P)$. Hence, if the correlation coefficient ρ_{imp} implied by (13) and the one actually observed in the period after t_0 are significantly different from each other, then ω must have changed.

In the lower part of table 4 we first present the ratio of the standard deviations of the common (Δy_t^{*P}) and idiosyncratic $(\widehat{y}_t^P)$ factors for the three subperiods. After 1980, this ratio seems to have fallen considerably, from 0.38 in 1960-80 to 0.33 in 1980-2000 to 0.14 in 1990-2000. Recall from our previous results that the variance of the country-specific trend growth factor has actually fallen itself. The results here imply that the volatility of the global trend growth factor has declined even more.

The last row of table 4 gives the implied correlations, ρ_{imp} for 1980-2000 and 1990-2000.[16] In view of the downward trend in the variance ratio between global and country-specific factors, it is not surprising that we find ρ_{imp} declines over time. If the entire post-1980 period is considered, the implied and actual correlations appear still quite similar. However, they clearly start to diverge after 1990: whereas the implied correlation drops markedly there is almost no decline in the correlation obtained from actual post-1990 data. In the context of our model here, this can only be explained through an offsetting increase in ω.

The stylized fact that the variability of global output shocks has decreased during the globalization period is now well-documented (see Kose, Otrok and Whiteman (2003, 2005) and Heathcote and Perri (2004)). Absent financial globalization, these changes in world business cycles should actually have lowered consumption correlations. The fact that consumption correlations have remained roughly constant after 1980 is therefore quite in line with the theoretical proposition that financial globalization should increase consumption correlations *ceteris paribus*.

6 Conclusion

Consumption based measures of international risk sharing seem to have remained unaffected by more than two decades of financial globalization. Whereas economic theory would suggest that the country-specific component of consumption volatility (conditional on output) should decrease and international consumption correlations should increase with financial integration, both measures have remained approximately constant after 1980

[16] These implied correlations are strictly valid only if Δy^{*P} and $\widehat{y}^{P}$ are orthogonal, an assumption that we find reasonable to maintain for the average country and that is confirmed in the data.

and throughout the 1990s. In this paper, we have put forward an explanation of this puzzle by arguing that the global decline in the volatility of output has almost offset the impact of financial globalization on consumption.

In our analysis, the distinction between permanent and transitory fluctuations in output and income is central, because consumption reacts primarily to permanent shocks and because it is plausibly much harder for countries to insure against permanent than against transitory fluctuations. Based on OECD data from 1960-2000, we illustrate that countries have become more insured in particular against permanent idiosyncratic shocks, in line with the ever better integration of financial markets.

We identify two reasons why basic measures of risk sharing have not picked up this trend, both of which are related to the by now well-researched global decline in business cycle volatility: first, the country-specific component of business cycles has become less volatile for the average OECD economy. We have shown that much of this drop in volatility is due to a more gradual response of output to permanent shocks. Since country-specific consumption growth reacts primarily to permanent idiosyncratic shocks, its volatility conditional on *current* country-specific output growth increases. This instance of Deaton's (1992) paradox will induce an upward bias in risk sharing regressions, offsetting the effects of financial globalization.

Secondly, the global component in permanent output fluctuations has also become less volatile which implies that international consumption correlations should have fallen *ceteris paribus*. The fact that they have stayed approximately constant is therefore consistent with the view that financial integration tends to increase consumption correlations.

Virtually all indicators of financial integration tell us that financial glob-

alization is happening. So far, consumption-based measures were an exception to this rule. This is troubling, because the very *raison d'être* of financial markets is the allocation of consumption risk and any assessment of the welfare benefits of financial integration will ultimately amount to a statement about consumption. It is therefore crucial to identify the factors that may have led international consumption comovements to defy the predictions of a wide class of theoretical models. To our knowledge, the results in this paper provide the first account of these factors: once we control for changes in international business cycles, consumption growth rates have become more correlated and relative consumption growth has become more independent from relative output growth – just as theory would predict.

We emphasize that this finding has nothing to say about whether the changes in international business cycles that we and others have identified over the globalization period could not by themselves be the outcome of financial or trade integration. Clearly, there are good theoretical reasons to believe that this may indeed be the case.

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Appendix: Construction of permanent GDP-values

We can rewrite the definition of $Y_t^P = (1-R)\sum_{k=0}^{\infty} R^k \mathbf{E}\left(Y_{t+k}\right)$ as

$$Y_t^P = (1-R)Y_t\left[1+\sum_{k=1}^{\infty} R^k \mathbf{E}\left(1+\frac{Y_{t+k}-Y_t}{Y_t}\right)\right]$$

and we can think of $\sum_{k=1}^{\infty} R^k \mathbf{E}\left(1+\frac{Y_{t+k}-Y_t}{Y_t}\right)$ as the percentage of total output that is transitory and can therefore be associated with the business cycle component. Since business cycles account for a small percentage of the level of GDP, we can use $\frac{Y_{t+k}-Y_t}{Y_t} \approx \log Y_{t+k} - \log Y_t$ to obtain the log-linear approximation:

$$\begin{aligned} Y_t^P &\approx (1-R)Y_t\left[1+\sum_{k=1}^{\infty} R^k \mathbf{E}\left(1+\log Y_{t+k}-\log Y_t\right)\right] \\ &= (1-R)Y_t\left[1+\frac{1}{1-R}-1+\sum_{k=1}^{\infty} R^k \mathbf{E}\left(\log Y_{t+k}-\log Y_t\right)\right] \\ &= (1-R)Y_t\left[\frac{1}{1-R}+\sum_{k=1}^{\infty} R^k \sum_{l=1}^{k} \mathbf{E}\left(\Delta y_{t+k+l}\right)\right] \end{aligned}$$

where $y_t = \log Y_t$.

Using that

$$\sum_{k=1}^{\infty} R^k \sum_{l=1}^{k} \Delta y_{t+k+l} = \frac{1}{1-R}\sum_{k=1}^{\infty} R^k \Delta y_{t+k}$$

we obtain

$$\begin{aligned} Y_t^P &\approx (1-R)Y_t\left[\frac{1}{1-R}+\frac{1}{1-R}\sum_{k=1}^{\infty} R^k \mathbf{E}\left(\Delta y_{t+k}\right)\right] \\ &= Y_t\left[1+\sum_{k=1}^{\infty} R^k \mathbf{E}\left(\Delta y_{t+k}\right)\right] \end{aligned}$$

and therefore with $y_t = \log Y_t$

$$y_t^P = \log\left[Y_t^P\right] \approx y_t + \sum_{k=1}^{\infty} R^k \mathbf{E}\left(\Delta y_{t+k}\right) = (1-R)\sum_{k=1}^{\infty} R^k \mathbf{E}\left(y_{t+k}\right) = y_t^P$$

We follow Crucini (1999) in comparing two alternative specifications for the permanent components of home and foreign output. First, we consider a univariate AR(1) process in growth rates of home and foreign output.

$$\Delta y_t^k = \rho_k \Delta y_{kt-1} + v_{kt} \text{ and } \Delta y_t^* = \rho^* \Delta y_{t-1}^* + v_t^*$$

While this specification implicitly assumes that there are no spillovers between home and foreign output, we also consider a VAR specification in output growth rates. Under the null of our model, aggregate consumption should be a sufficient statistic for expected future levels of output. We therefore follow the methodology by Campbell and Shiller (1989) and, besides home and foreign output growth, also include (relative) consumption as an endogenous state. We then estimate the VAR

$$\mathbf{\Delta x}_t = \mathbf{A\Delta x}_{t-1} + \boldsymbol{\varepsilon}_t$$

where $\mathbf{x}_t = \left[\begin{array}{ccc} y_{kt} & y_t^* & c_k - c_t^* \end{array} \right]'$ and $\mathbf{A}$ is the 3×3 coefficient matrix.

Using the approximation derived we write the relative permanent levels of output as

$$y_t^P - y_t^{*P} \approx y_{kt} - y_t^* + \mathbf{E}_t \left\{ \sum_{l=1}^{\infty} \frac{\Delta y_{t+l}^k - \Delta y_{t+l}^*}{(1+r)^k} \right\}$$

To construct the relative permanent values $y^p - y^{*P}$ from the VAR-process, we use the Hansen-Sargent prediction formula to get

$$\mathbf{E}_t \left\{ \sum_{l=1}^{\infty} \frac{\Delta y_{t+l}^k - \Delta y_{t+l}^*}{(1+r)^k} \right\} = \mathbf{h}' \left[\frac{\mathbf{A}}{1+r} \right] \left[\mathbf{I} - \frac{1}{1+r} \mathbf{A} \right]^{-1} \mathbf{\Delta x}_t$$

where $\mathbf{h}' = \left[\begin{array}{ccc} 1 & -1 & 0 \end{array} \right]$. In the case of the AR(1)-process we obtain a similar expression in which $\mathbf{A} = \left[\begin{array}{cc} \rho_k & 0 \\ 0 & \rho^* \end{array} \right]$.

The results reported in the main text, are based on values of $y^P - y^{*P}$ constructed with the real interest rate, r, set to 0.02. We provide the results from a battery of robustness checks in an extended version of this appendix (available from the authors upon request), where we allow r to vary across a wide range of different values and also across countries. Our results remain unchanged by these exercises.

Table 1: Basic risk sharing regressions

United States	OECD		
1960-90	1960-1980	1980-2000	1990-2000
	Estimate of b		
0.16	0.79	0.89	0.85
(4.6)	(19.45)	(10.22)	(6.27)

NOTES: Panel regressions of the form $\Delta c_t^k - \Delta c_t^* = b(\Delta y_t^k - \Delta y_t^*) + \delta + \mu^k + \varepsilon_t^k$ where δ, μ^k and ε_t^k are the constant term, the country (region) fixed effect and the residual term respectively. Numbers in parentheses are t-statistics based on the weighted least squares procedure described in the main text.

Table 2: Sharing of permanent and transitory idiosyncratic risk

Specification	United States	OECD		
for $\Delta y - \Delta y^*$	(1960-90)	1960-80	1980-2000	1990-2000
		Permanent component: b_P		
AR(1)	0.08	0.57	0.42	0.30
	(3.18)	(15.53)	(7.42)	(2.71)
VAR(1)	0.04	0.58	0.43	0.19
	(1.52)	(12.53)	(5.84)	(1.66)
		Transitory component: b_T		
AR(1)	0.01	-0.40	-0.28	0.07
	(0.08)	(-3.69)	(-2.93)	(0.41)
VAR(1)	0.16	0.26	0.14	0.20
	(3.23)	(3.13)	(1.33)	(1.54)

NOTES: Panel regressions of the form $\Delta c_t^k - \Delta c_t^* = b_P(\Delta y_t^{kP} - \Delta y_t^{*P}) + \delta + \mu^k + \varepsilon_t^k$ and $\Delta c_t^k - \Delta c_t^* = b_T(\Delta y_t^{kT} - \Delta y_t^{*T}) + \delta + \mu^k + \varepsilon_t^k$ where δ, μ^k and ε_t^k are the constant term, the country (region) fixed effect and the residual term respectively. Numbers in parentheses are t-statistics based on the weighted least squares procedure described in the main text. The first column identifies which specification for $\Delta y - \Delta y^*$ was used in the construction of permanent and transitory components.

Table 3: Business cycle volatility and risk sharing regressions

	Business cycle volatility $var(\widehat{y})^{1/2}$	Trend volatility $var(\widehat{y}^P)^{1/2}$	Long-Run variance ratio $\phi = \frac{var(\widehat{y}^P)}{var(\widehat{y})}$	Implied degree of globalization $b_{adj} = b/\phi$
OECD				
1960-80	0.028	0.030	1.14 (0.03)	0.70 (0.03)
1980-2000	0.020	0.027	1.80 (0.06)	0.50 (0.02)
1990-2000	0.017	0.025	2.16 (0.22)	0.39 (0.05)
United States				
1960-90	0.034	0.042	1.56 (0.02)	0.10 (0.005)

NOTES: The table presents the impact of changes in business cycle and long-run volatility ($var(\widehat{y})$ and $var(\widehat{y}^P)$ respectively) on the risk sharing coefficient. Calculations of permanent components $\widehat{y}^P$ based on the VAR. For the calculation of $b_{adj} = b/\phi$ the estimates of b from table 1 were used. Standard errors(in parentheses) reported were obtained as follows: we repeatedly estimated the panel, dropping the observations for one country or region at a time. The respective statistics was then re-calculated and stored. This procedure gives us 22 (or 50 for the U.S. without Washington D.C.) different realizations of each the three parameters. The standard errors reported are the standard deviations over these 22 (50) realizations. For further details see section 5.2.

Table 4: Int'l consumption correlations and financial globalization

	United States	OECD		
	1960-90	1960-1980	1980-2000	1990-2000
	Data			
$corr(\Delta c, \Delta c^*)$	0.54	0.35	0.34	0.31
$corr(\Delta y, \Delta y^*)$	0.78	0.47	0.52	0.60
	Variability of global relative to country-specific component and implied international consumption correlations			
$\frac{var(\Delta y^{*P})}{var(\widehat{y}^P)}$		0.38	0.33	0.14
ρ_{imp}		—	0.33	0.23

NOTES: Reported correlations are the cross-sectional (across countries or US states) averages of the correlations of consumption/output growth rates with the respective world aggregate. $\rho_{imp} = \rho\left((1+\delta)\left(1+\delta\rho^2\right)^{-1}\right)^{1/2}$, where ρ is the correlation between consumption growth rates from the preceding subperiod. and δ is the percentage change between successive periods in the ratio $var(\Delta y^{*P})/var(\widehat{y}^P)$.

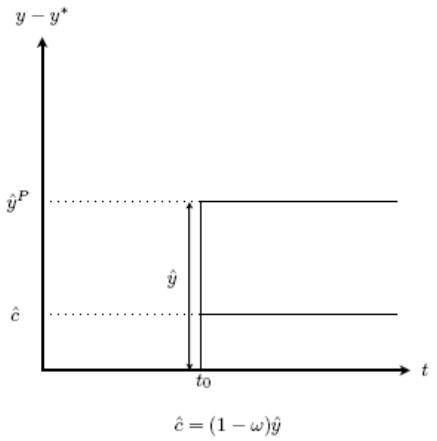

Figure 1: Dynamic response of output to a permanent shock. Case 1: immediate adjustment, $\phi = 1$.

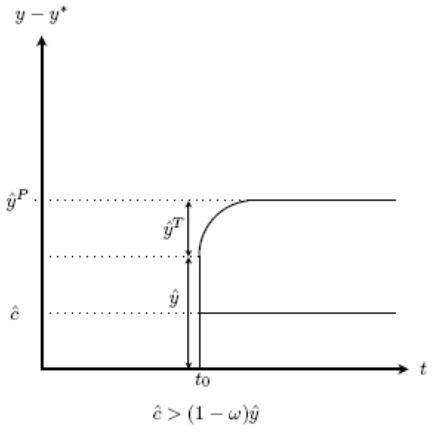

Figure 2: Dynamic response of output to a permanent shock. Case 2: gradual adjustment, $\phi > 1$.