Common Correlated Effects and International Risk Sharing*†

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Abstract

Existing studies of international risk sharing rely on the highly restrictive assumption that all economies are characterized by symmetric preferences and uniform transmission of global shocks. We relax these homogeneity constraints by modeling aggregate and idiosyncratic fluctuations as unobserved components, and we use Pesaran's (2006) common correlated effects estimator to control for common factors and cross-sectional heterogeneity. We compare the proposed approach with the conventional ones using data from Penn World Table 9.0 for 120 countries. While we do not detect a significant increase in risk sharing during the last four decades, our results confirm that consumption is only partially smoothed internationally and risk sharing is directly related to the level of development.

Keywords: International risk sharing, Consumption insurance, Panel data, Cross-sectional dependence, Heterogeneous effects
JEL codes: C23, C51, E21, F36

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1 Introduction

Since the early contributions by Cochrane (1991), Mace (1991), and Obstfeld (1994), a number of consumption risk sharing tests have been presented in the literature. Published research finds excess sensitivity of consumption to income shocks and this has been interpreted as lack of international risk sharing. To maintain analytical tractability, the derivation and implementation of risk sharing tests usually relies on several homogeneity assumptions that are unlikely to hold in worldwide panels: all economies are assumed to be characterized by symmetric preferences and uniform transmission of global shocks. The extent of risk sharing is then estimated by a panel data regression of cross-sectionally demeaned consumption on cross-sectionally demeaned income. However, if the homogeneity assumptions underlying the analysis are violated, the results may be biased.

We extend the existing literature by taking into account various sources of heterogeneity. Specifically, we allow for cross-country variation 1) in preferences, 2) in the transmission of global income shocks, and 3) in the sensitivity of consumption to income shocks. Considering these sources of heterogeneity is worthwhile for the following reasons:

1. The underlying theory suggests that if a country has full access to international risk sharing opportunities, consumption will be independent of idiosyncratic income shocks. However, this does not necessarily imply uniform consumption growth around the world. Country-level and global consumption will move in lockstep if preferences are symmetric across countries, but they will diverge if risk aversion or discount factors are heterogeneous (Obstfeld, 1989, 1994). Hence, even under perfect risk sharing, consumption paths can differ from each other due to heterogeneous preferences.

2. If risk sharing is imperfect, consumption will also be affected by idiosyncratic income fluctuations. This in turn raises the question of how to isolate idiosyncratic income shocks from global ones. Due to differences in their productive and financial structure, regulations, and their participation in international trade, countries may be affected by aggregate shocks to varying degrees. For example, a country with a disproportionally large export sector may face greater income fluctuations caused by aggregate sources than a country that does not participate in international trade. Accordingly, idiosyncratic shocks can be obtained by controlling the extent of global shocks transmitted to individual countries; an extent that is unlikely to be uniform throughout the world (Giannone and Lenza, 2010).

3. Finally, because of differences in the quality of smoothing channels, the effects of idiosyncratic income shocks on consumption may also vary across countries.

By taking into account the aforementioned sources of cross-sectional heterogeneity, we can more accurately estimate the extent of international risk sharing.

We argue that the appropriate method for filtering out the unobserved common factors from the observed variables should allow for the heterogeneity of countries in terms of their preferences and exposure to aggregate risk. Consequently, and at odds with the existing literature, we let global factors have country specific loading coefficients. In addition, we relax the homogeneity assumption behind pooled or fixed effects estimation and employ a mean-group type estimator that is robust to heterogeneous country characteristics. Due to these refinements, the proposed approach is better at isolating idiosyncratic fluctuations and less susceptible to bias than the cross-sectional demeaning method. We illustrate the performance of the considered methods under cross-sectional dependence and heterogeneity by analyzing a wide range of countries.

Accounting for these features of the data is consequential as some homogeneity restrictions have a significant impact on the estimates.

We contribute to the existing literature in several important ways. First, we highlight the shortcomings of the conventional approach to analyze international risk sharing. Second, we re-evaluate earlier results on the lack of perfect risk sharing using a more flexible econometric model that isolates idiosyncratic fluctuations in the data. Specifically, we are the first ones to apply the common correlated effects (CCE) estimator of Pesaran (2006) to the analysis of international risk sharing. Third, while earlier studies have focused on smaller more homogeneous sets of economies, our large panel of 120 countries allows us to analyze risk sharing along a variety of economic characteristics. Fourth, we look at the change in the extent of risk sharing over the past forty years, but find no evidence to support the notion that financial globalization has led to an increase in international consumption smoothing.

2 The Conventional Approach

Regression based risk sharing (or consumption insurance) tests are based on the null hypothesis of market completeness, or the possibility to redistribute wealth (hence, consumption) across all date-event pairs. Under market completeness, the solution to the representative agent's maximization problem ensures that marginal utility growth is equalized across agents and depends on aggregate factors but not on individual shocks (Cochrane, 1991; Mace, 1991; Obstfeld, 1994). Assuming CRRA utility functions, the risk sharing hypothesis can be tested using the following equation

$$c_{it} = \alpha_i + \gamma_i^c \bar{c}_t + \beta_i x_{it} + \varepsilon_{it} \ , \ i = 1 \ldots N, \ t = 1 \ldots T \ , \tag{1}$$

where c_{it} is a consumption measure for country i, $\bar{c}_t$ is an aggregate measure of consumption, and x_{it} is an idiosyncratic variable. Market completeness implies $\gamma_i^c > 0$ and $\beta_i = 0$. If the discount factors and the coefficients of relative risk aversion are assumed to be equal across countries, the coefficients γ_i^c can be shown to take a unit value. However, such homogeneity is unlikely in reality: Obstfeld (1989) found some evidence against the hypothesis of $\gamma_i^c = 1$ even in countries with similar characteristics, such as Germany, Japan and the United States. Nevertheless, to maintain tractability of the analysis, many papers in the field have built on these homogeneity assumptions, under which the test equation becomes

$$c_{it} - \bar{c}_t = \alpha_i + \beta_i x_{it} + \varepsilon_{it} \ . \tag{2}$$

The consumption risk sharing test is based on the null hypothesis $H_0 : \beta_i = 0$, where β_i can be regarded as the extent of the departure from perfect risk sharing. For example, Asdrubali et al. (1996); Crucini (1999); Crucini and Hess (2000); Grimard (1997); Jalan and Ravallion (1999), and many others in the last decade noted that the relative size of the estimated slope coefficient can be interpreted as a measure of the degree of insurance or risk pooling. The rejection of the null hypothesis implies that agents do not use an insurance mechanism to fully offset idiosyncratic shocks to their endowments, which are consequently transmitted to consumption.

In virtually all macroeconomic implementations of equation (2), the variable x_{it} containing idiosyncratic shocks is replaced by a proxy for idiosyncratic income, which in turn is calculated as a difference between the individual country's income and a measure of aggregate income. With these modifications the tested relationship becomes

$$c_{it} - \bar{c}_t = \alpha_i + \beta_i (y_{it} - \bar{y}_t) + \varepsilon_{it} \ , \tag{3}$$

where y_{it} is an income measure for country i, and $\bar{y}_t$ is a measure of aggregate income. To obtain an overall β coefficient for the analyzed set of countries, most researchers pool the data and estimate the fixed effects regression

$$c_{it} - \bar{c}_t = \alpha_i + \beta(y_{it} - \bar{y}_t) + \varepsilon_{it} \ , \tag{4}$$

which imposes an additional layer of homogeneity on the model.

Equation (4) is the basis for several recent influential empirical studies, such as Sorensen and Yosha (2000), Sorensen et al. (2007), and Kose et al. (2009) among others. In these studies, the consumption and income measures entering the analysis are consumption growth and real gross domestic product (GDP) growth, respectively. Correspondingly, β is interpreted as the effect of idiosyncratic real GDP growth on idiosyncratic consumption growth. If the aggregates, $\bar{c}_t$ and $\bar{y}_t$, are cross-sectional means, then the differencing operations in equation (4) will produce cross-sectionally demeaned variables. Other studies, for example Asdrubali et al. (1996), Lewis (1997), Sorensen and Yosha (1998), and Fratzscher and Imbs (2009), replace the explicit cross-sectional demeaning in equation (4) with an implicit one by including a time dummy d_t in the fixed effects regression

$$c_{it} = \alpha_i + d_t + \beta y_{it} + \varepsilon_{it} \ . \tag{5}$$

Artis and Hoffmann (2006) derive equation (4) by relying on a different theoretical framework proposed by Crucini (1999). They model country specific income, y_{it}, as a mixture of the level of pooled real GDP in participating countries and the level of domestic real GDP. Their results rely on an assumption of perfect symmetry, where each country is assumed to pool the same proportion of its income. However, similarly to the assumption of equal discount factors and coefficients of risk aversion across countries in the classical framework, this assumption is also likely overly restrictive when the analysis is carried out with a heterogeneous set of economies.

It is important to remember, that the correlation between consumption and income is also a measure of intertemporal consumption smoothing (see for example Ostergaard et al., 2002). Asdrubali and Kim (2008) and Ho et al. (2010) analyzed risk sharing and intertemporal smoothing jointly

$$c_{it} = \alpha_i + (1-\gamma)(1-\omega)y_{it} + \gamma\bar{c}_t + \varepsilon_{it} \ , \tag{6}$$

where γ captures the extent of risk sharing and ω captures the extent of intertemporal smoothing, while $\beta = (1-\gamma)(1-\omega)$ still reveals the impact of income shocks on consumption.[1] Notice that equation (6) does not contain pooled income, $\bar{y}_t$, as a control variable, and all coefficients are "global", without an i subscript. Again, both of these features of the model can be attributed to the unrealistic assumption of identical countries, specifically that they pool equal shares of their incomes and use equal fractions of their disposable incomes to smooth consumption.

Similarly, a hybrid model, in which the heterogeneous impact of income shocks is neglected,

$$c_{it} = \alpha_i + \beta_i(y_{it} - \bar{y}_t) + \gamma_i\bar{c}_t + \varepsilon_{it} \ , \tag{7}$$

discards the arguments made in bullet point 2 of Section 1. When two, otherwise equivalent, countries are participating in international trade to a different extent, the same global shock will affect the two countries to a different degree. However, this does not mean that the transmitted shocks, albeit having different magnitudes, should be considered idiosyncratic. Strictly speaking, the shocks affecting the two countries

[1] In line with most of the early literature, we refer to β as a measure of risk sharing, but in the framework proposed by Asdrubali and Kim (2008) β can be interpreted as a joint measure of risk sharing and intertemporal smoothing.

are triggered by the same underlying factor, or global shock, and not allowing factor loadings to vary by country will make it impossible to eliminate the transmitted amount *of the global shock* from individual country incomes. Consequently, $y_{it} - \bar{y}_t$ does not produce idiosyncratic income fluctuations.

3 An Alternative Approach

We propose to deal with the cross-sectional variation in country characteristics and the estimation of idiosyncratic effects by taking advantage of an unobserved component model (Harvey, 1989). Although neither aggregate nor idiosyncratic shocks are directly measured, a particular country's observed income, y_{it}, can be decomposed into two analogous unobserved components. By definition, pooled income will follow global cycles that can be modeled by common factors, $\boldsymbol{f}_t$, and its contribution to a particular country's observed income can be captured by the factor loadings, $\boldsymbol{\lambda}_{i,y}$,

$$y_{it} = \boldsymbol{\lambda}'_{i,y}\boldsymbol{f}_t + \xi^y_{it} \ , \tag{8}$$

where $\boldsymbol{\lambda}_{i,y}$ allows countries to be heterogeneous in terms of their sensitivity to global shocks. The term $\boldsymbol{\lambda}'_{i,y}\boldsymbol{f}_t$ yields the amount of fully diversified income for country i, and the balance, $\xi^y_{it} = y_{it} - \boldsymbol{\lambda}'_{i,y}\boldsymbol{f}_t$, is the idiosyncratic income. Applying a similar logic to the calculation of idiosyncratic consumption, and approximating the common factors with cross-sectional means of the variables, we obtain the more general model

$$c_{it} - \gamma^c_i \bar{c}_t = \alpha_i + \beta_i(y_{it} - \tilde{\gamma}^y_i \bar{y}_t) + \varepsilon_{it} \ , \tag{9}$$

or

$$c_{it} = \alpha_i + \beta_i y_{it} + \gamma^c_i \bar{c}_t + \gamma^y_i \bar{y}_t + \varepsilon_{it} \ , \tag{10}$$

where the β_i coefficient measures the extent to which idiosyncratic shocks to income are channeled into idiosyncratic consumption. The approximation of common factors by cross-sectional averages is advantageous for two reasons: 1) the analysis of risk sharing focuses on consistent estimation of the β coefficient, but it does not concern itself with common factors per se, and 2) Westerlund and Urbain (2015) have shown that this approximation results in lower bias of the β estimate than competing approaches based on direct estimation of the common factors. The country specific $\gamma_i^y = -\beta_i \tilde{\gamma}_i^y$ and γ_i^c coefficients allow the amount of income and consumption driven by global shocks to vary across economies. A more detailed discussion of this model follows in Section 4.

4 Empirical Strategy

The international risk sharing hypothesis postulates that consumption across countries follows a similar pattern, and idiosyncratic deviations from this pattern cannot be predicted by idiosyncratic explanatory variables. The presence of a similar pattern across countries can be tested by the cross-sectional dependence (CD) statistic of Pesaran (2004). This test is based on the pairwise correlation of the cross-sectional units, and has been shown to have good finite sample properties in heterogeneous panels. If the null hypothesis of cross-sectional independence is rejected, the co-movement of variables across countries may be modeled by common factors, and idiosyncratic components can be obtained by an orthogonal projection of the data onto the common factors. A relationship between idiosyncratic consumption and income can then be estimated and tested for significance.

Pesaran's (2006) common correlated effects (CCE) estimator, which he proposed to deal with dependencies across units in heterogeneous panels, is an ideal tool for estimating β_i, the effect of idiosyncratic income on idiosyncratic consumption. The

CCE estimator lends itself to this task because it accounts for common factors, such as global cycles, allows for individual specific effects of these factors, and produces coefficient estimates based on idiosyncratic fluctuations in the data.

The CCE estimator is equivalent to ordinary least squares applied to an auxiliary regression augmented with the cross-sectional means of the variables. In other words, the β_i^{CCE} estimatesare identical to ordinary least squares estimates of β_i in our proposed model (10). The CCE estimator partitions the regression in (10) by projecting consumption and income orthogonally with respect to their cross-sectional means. The estimation can also be viewed as a two stage regression. In the first stage, the common effects are filtered out from the data by regressing each variable on the cross-sectional averages of all variables in the model

$$c_{it} = a_{i,c} + \lambda_{i,c}^{c}\bar{c}_t + \lambda_{i,c}^{y}\bar{y}_t + \xi_{it}^{c} \ , \tag{11}$$

$$y_{it} = a_{i,y} + \lambda_{i,y}^{c}\bar{c}_t + \lambda_{i,y}^{y}\bar{y}_t + \xi_{it}^{y} \ . \tag{12}$$

In the second stage, the CCE estimate of an individual β_i is obtained by regressing the residual $\hat{\xi}_{it}^{c}$, capturing idiosyncratic consumption, on the residual $\hat{\xi}_{it}^{y}$, capturing idiosyncratic income. While the λ coefficients in (11) and (12) can not be meaningfully interpreted (see Pesaran, 2006; Westerlund and Urbain, 2015), the residuals $\hat{\xi}_{it}^{c}$ and $\hat{\xi}_{it}^{y}$ are valid estimates of the idiosyncratic components and can be compared to cross-sectionally demeaned consumption and income. Note that the latter may not be free of aggregate shocks: if the effect of global cycles differs across countries, cross-sectional demeaning will not be able to isolate the idiosyncratic variation in the data and will therefore lead to biased conclusions about the extent of risk sharing.

Most empirical analyses focus on testing the risk sharing hypothesis with differenced data. However, several recent studies, including Becker and Hoffmann (2006) and Artis and Hoffmann (2012), have examined the implications of risk sharing in the long run by exploiting the information contained in the levels of the variables. Conveniently, our proposed estimation procedure does not depend on the transformation of the variables: Kapetanios et al. (2011) proved that the CCE estimators are consistent as long as the regression residuals are stationary. The rejection of a unit root in ε_{it} (in equation 10) implies that c_{it}, y_{it}, and $\boldsymbol{f}_t$ are cointegrated, and additional information can be obtained about risk sharing within an error correction model (Fuleky et al., 2015; Leibrecht and Scharler, 2008; Pierucci and Ventura, 2010).

For an individual country, the deviation from the long run equilibrium relationship between idiosyncratic income and consumption, after controlling for permanent global shocks, is captured by the residual, $\hat{\varepsilon}_{it}$, in equation (10). The speed, κ, at which this equilibrium error is corrected can then be estimated along with the extent of risk sharing in the short run, β_i^{SR}, in the following error-correction model

$$\Delta c_{it} - \gamma_i^{c,SR}\overline{\Delta c}_t = \alpha_i^{SR} + \kappa \hat{\varepsilon}_{it}^{LR} + \beta_i^{SR}(\Delta y_{it} - \tilde{\gamma}_i^{y,SR}\overline{\Delta y}_t) + \varepsilon_{it}^{SR} \ , \tag{13}$$

or

$$\Delta c_{it} = \alpha_i^{SR} + \kappa \hat{\varepsilon}_{it}^{LR} + \beta_i^{SR}\Delta y_{it} + \gamma_i^{c,SR}\overline{\Delta c}_t + \gamma_i^{y,SR}\overline{\Delta y}_t + \varepsilon_{it}^{SR} \ , \tag{14}$$

where $\hat{\varepsilon}_{it}^{LR} = c_{it} - \hat{\alpha}_i^{LR} - \hat{\beta}_i^{LR} y_{it} - \hat{\gamma}_i^{c,LR}\bar{c}_t - \hat{\gamma}_i^{y,LR}\bar{y}_t$. Here, the heterogeneous impact of transitory global shocks is filtered out by including in the regression the cross-sectional means of differenced consumption and income, $\overline{\Delta c}_t$ and $\overline{\Delta y}_t$, with country specific coefficients, $\gamma_i^{c,SR}$ and $\gamma_i^{y,SR}$, respectively (see also Holly et al., 2010, Sec. 5.4). The cross-sectional averages also control for potential endogeneity bias arising due to common factors in differenced income and the error term. The dynamic specification of the error-correction model follows the recommendation of Davidson et al. (1978), who find that persistence in annual data can largely be captured by the first lag.

Under a random coefficient model, the simple averages of the individual CCE estimators of β_i^{LR} and β_i^{SR} are consistent estimators of the overall β^{LR} and β^{SR}, respectively. These mean-group estimators are defined as

$$\hat{\beta}^{LR} = \frac{1}{N}\sum_{i=1}^{N}\hat{\beta}_i^{LR} \quad \text{and} \quad \hat{\beta}^{SR} = \frac{1}{N}\sum_{i=1}^{N}\hat{\beta}_i^{SR}\ . \tag{15}$$

Apart from results for the fixed effects model in equation (3), all other results in our study are obtained by mean-group type aggregation of individual estimates. Coakley et al. (2001) showed that, in contrast to pooled and fixed effects estimators, mean-group estimators are robust to dependence between the coefficients and the regressors along the cross-sectional dimension. Furthermore, Coakley et al. (2006) found that among a variety of mean-group estimators, including one based on a cross-sectionally demeaned regression specified in equation (4), the CCE mean-group estimator is the most robust to general settings, such as regressors and errors sharing common factors with possibly correlated factor loadings.

5 Data and Results

Our analysis is based on annual data obtained from the Penn World Table 9.0, released in June 2016 (Feenstra et al., 2015). This is a comprehensive dataset, covering more than 170 countries over a fairly long time span. We use the subperiod 1970 - 2014, which yields 120 countries with population over one million and continuously available annual data. The analysis of such a large heterogeneous panel is a distinguishing feature of our study; the existing literature focuses on smaller sets of rather homogeneous countries.

From the Penn World Tables we use purchasing power parity converted GDP per capita and consumption per capita at 2005 constant prices.[2] The analyzed series are

[2]The basic risk sharing equation can be augmented by additional regressors, such as proxies for

Table 1: Diagnostic Tests for Individual Variables

	Levels		Differences	
	$\log C$	$\log Y$	$\Delta \log C$	$\Delta \log Y$
CD	268.92*	232.58*	19.81*	40.76*
$CIPS_{\mu}$	-1.49	-1.45	-2.78*	-2.37*
$CIPS_{\mu,t}$	-2.41	-2.26	-3.14*	-2.76*

Note: Pesaran's (2004) cross-sectional independence test statistic (CD) follows a standard normal distribtion. The lag length for Pesaran's (2007) panel unit root test ($CIPS$) is set to $T^{1/3} \approx 4$. The 5 % critical values for $CIPS_{\mu}$ (the model includes an intercept) and $CIPS_{\mu,t}$ (the model includes an intercept and a linear trend) are -2.06 and -2.55, respectively. Statistical significance at the 5% level is denoted by *.

comparable to those in other datasets, such as the World Bank's World Development Indicators. They are expressed in real terms in a common currency to make comparisons across countries and time feasible. Because these variables tend to exhibit exponential growth, we apply a logarithmic transformation to them in our analysis. The diagnostic statistics displayed in Table 1 indicate that log-consumption and log-income levels are cross-sectionally dependent and follow stochastic trends. The log-differenced series are also cross-sectionally dependent, but they do not contain unit roots.

Table 2 displays the results of diagnostic tests applied to the residuals of each model we consider in our analysis (see the note in the table for model specifications). In each regression, we test the residuals for cross-sectional dependence and, in long run models when the data are in log-levels, for non-stationarity. We use the CD statistic proposed by Pesaran (2004) for the former, and the $CIPS$ statistic of Pesaran (2007) for the latter (see also Banerjee and Carrion-i Silvestre, 2014; Holly et al., 2010). None of the traditional models relying on cross-sectional demeaning (FE, DEM, HYB) are able to control for common factors, or isolate idiosyncratic shocks, and the FE

financial development, net foreign income flows, etc. However, because we wanted to directly relate our study to the seminal contributions in the literature cited in Section 2, we did not include extra control variables in our analysis.

Table 2: Residual Diagnostic Tests for the Whole Sample

	Long Run				Short Run			
	FE	DEM	HYB	CCE	FE	DEM	HYB	CCE
CD	4.46*	11.02*	6.67*	-0.47	11.77*	14.36*	10.62*	0.62
$CIPS_{\mu}$	-1.31	-1.76	-2.53*	-2.71*	—	—	—	—

Note: See also the notes in Table 1. Models in columns: Long Run FE is the fixed effects regression in equation (4); Long Run DEM is the regression using cross-sectionally demeaned consumption and income in equation (3); Long Run HYB is a hybrid model between the DEM and CCE models specified in equation (7); Long Run CCE is the model in equation (10); Short Run FE and Short Run DEM are equivalent to their long run counterparts but evaluated with differenced data; Short Run HYB is an error correction model, or equation (7) with data in log-differences and augmented with the residuals from the Long Run HYB model; Short Run CCE is the error-correction model in equation (14).

Table 3: Chow Test of Poolability for Select Subsamples

	Long Run			Short Run		
	DEM	HYB	CCE	DEM	HYB	CCE
Whole Sample	31.15*	36.78*	32.12*	6.91*	4.23*	3.79*
OECD	27.02*	39.21*	30.39*	4.16*	2.41*	2.44*

Note: See also the notes in Tables 1 and 2. The Chow test for the poolability of the data is an F test of stability for the coefficients of a panel model. Rejection of the null hypothesis implies that the individual slope coefficients are not the same, $\beta_i \neq \beta$, and therefore pooled or fixed effects estimation is inappropriate.

and DEM regressions are spurious with unit roots in the residuals. Although the literature frequently resorts to pooled or fixed effects (FE) estimation of the model coefficients, the results in Table 3 indicate that imposing the same β for each country is inappropriate even in the case of relatively homogenous OECD countries.

5.1 "Idiosyncratic" Fluctuations

Due to their great diversity, the countries in our analysis vary in terms of their susceptibility to global shocks. The rejection of cross-sectional independence for the residuals of the FE model in equation (4), DEM model in equation (3), and the HYB model in equation (7) indicates that cross-sectional demeaning is not able to fully isolate the idiosyncratic fluctuations in the variables. In other words, the unit coefficients imposed on the aggregates do not reflect the true influence of global shocks on country level variables, and they give rise to residual common factors in the regression. If the countries were homogeneous in terms of risk aversion, time preference, and endowments, the global shocks would have a unit loading for each country, and cross-sectional demeaning would be an appropriate method to calculate the idiosyncratic components. However, when the countries are heterogeneous, and the impact of global shocks differs across countries, the first stage regressions (11) and (12) are more appropriate to estimate idiosyncratic variation.

The estimated degree of risk sharing will be biased if the global shocks are not fully filtered out from the variables because $\hat{\beta}_i$ will, at least in part, attribute aggregate fluctuations in consumption to aggregate fluctuations in income. Global factors are essentially lurking variables that confound the relationship between the regressor and the dependent variable. The problem gets exacerbated by the restriction placed on β in the FE model in equation (4), which pushes country specific effects of income shocks into the error term and further distorts the estimates due to the prevalence of unit roots in the residuals. Consequently, the diagnostic tests indicate that only models augmented by cross-sectional averages, that is, equations (10) and (14), yield statistically acceptable results.

To illustrate the disagreement between the two methods in our heterogeneous data set, we examine the correlation of the idiosyncratic components estimated by the first stage regressions and cross-sectional demeaning. Figure 1 shows the distribution of the correlation coefficients $Cor(\hat{\xi}^c_{it}, c_{it} - \bar{c}_t)$ and $Cor(\hat{\xi}^y_{it}, y_{it} - \bar{y}_t)$ when the data is in log-levels. The correlation is below 0.80 for over two thirds of the countries. The

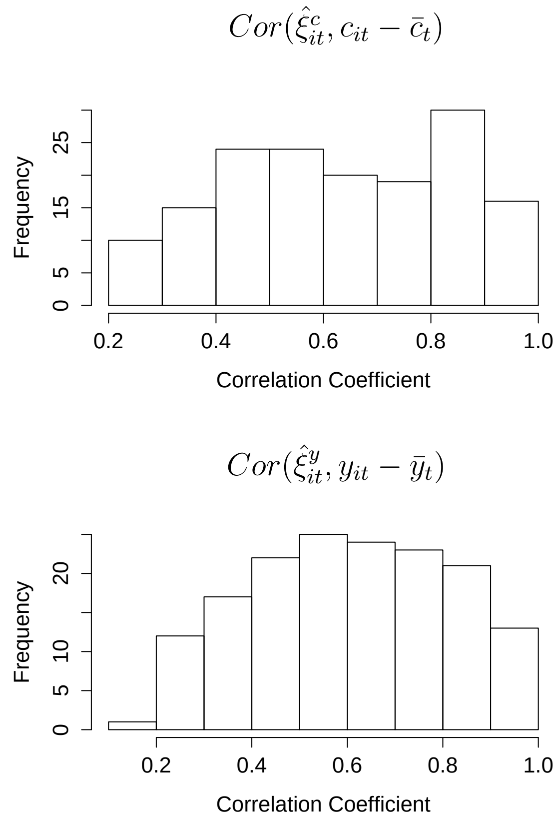

Figure 1: Distribution of correlation coefficients $Cor(\hat{\xi}^c_{it}, c_{it} - \bar{c}_t)$ and $Cor(\hat{\xi}^y_{it}, y_{it} - \bar{y}_t)$. The idiosyncratic components, $\hat{\xi}^c_{it}$ and $\hat{\xi}^y_{it}$, are estimated in (11) and (12), and the cross-sectionally demeaned variables, $c_{it} - \bar{c}_t$ and $y_{it} - \bar{y}_t$, appear directly in equations (3), (4), and (7). All analyzed series are in log-levels.

correlation is close to unity if the country specific income and consumption closely follow their aggregate counterparts, but close to zero when a country is not influenced by global shocks. In the former case both methods can successfully eliminate the global effects. However, in the latter case, $c_{it} - \bar{c}_t$ and $y_{it} - \bar{y}_t$ introduce mirror images of the global shocks into the demeaned variables, while factor loadings equal to zero ensure that $\hat{\xi}^c_{it}$ and $\hat{\xi}^y_{it}$ remain void of global shocks.

This discrepancy between the methods is highlighted using two representative countries in Figure 2. The plots illustrate the evolution of idiosyncratic components and demeaned variables, and it is evident that the latter are trending. Those stochastic trends are either introduced (Central African Republic - not sensitive to global shocks) or not fully removed (China - highly sensitive to global shocks) by cross-sectional demeaning. The stochastic trends show up on both the left and the right hand side of equation (3), which leads to bias in the individual β^{LR}_{DEM} estimates for two reasons. First, $\hat{\beta}^{LR}_{DEM}$ attributes the trend in demeaned consumption to the trend in demeaned income. Second, the diagnostic tests of the regression residuals in Table 2 imply that cross-sectionally demeaned income and consumption are not cointegrated, and the β^{LR}_{DEM} estimates are spurious. When the model in equation (3) is evaluated with log-differenced series, the β^{SR}_{DEM} estimates do not suffer from the issues related to non-stationarity, but they are influenced by the lingering aggregate effects in the cross-sectionally demeaned data. These illustrations further corroborate our earlier finding that imposing a unit loading coefficient on the aggregates leaves the demeaned regression misspecified and incapable of filtering out the common factors from our heterogeneous panels.

5.2 Estimates of Risk Sharing Across Methods

We now turn to the discussion of the statistically defensible CCE coefficient estimates, which we also contrast with the estimates obtained by conventional methods. Table 4

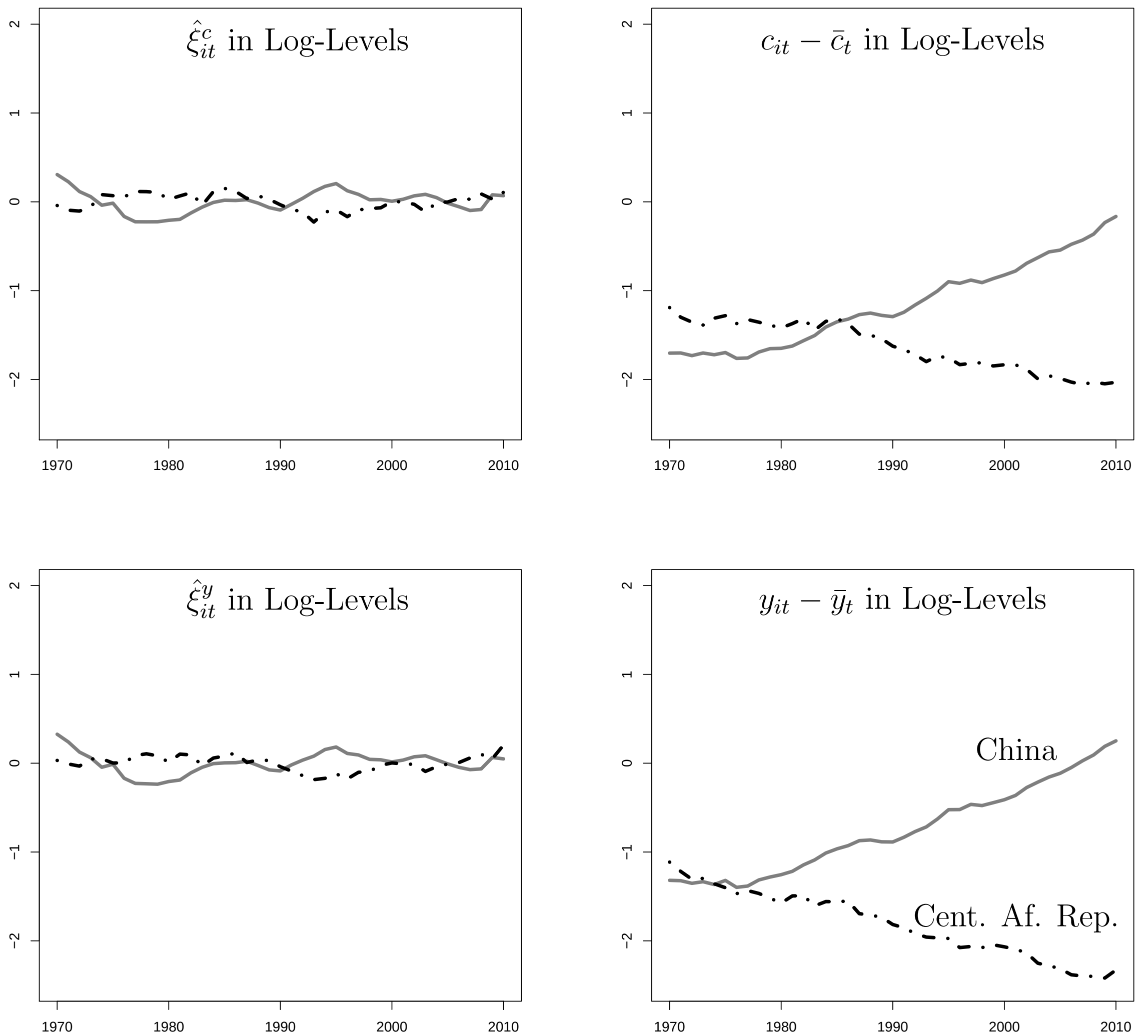

Figure 2: Plots of idiosyncratic components, $\hat{\xi}^c_{it}$ and $\hat{\xi}^y_{it}$, and cross-sectionally demeaned variables, $c_{it} - \bar{c}_t$ and $y_{it} - \bar{y}_t$ for two representative countries: China (solid line) highly sensitive to global shocks and Central African Republic (dash-dotted line) not sensitive to global shocks. All analyzed series are in log-levels.

Table 4: Coefficient Estimates

Country Group	$\hat{\beta}^{LR}_{FE}$	$\pm$	$\hat{\beta}^{LR}_{DEM}$	$\pm$	$\hat{\beta}^{LR}_{HYB}$	$\pm$	$\hat{\beta}^{LR}_{CCE}$	$\hat{\beta}^{SR}_{FE}$	$\pm$	$\hat{\beta}^{SR}_{DEM}$	$\pm$	$\hat{\beta}^{SR}_{HYB}$	$\pm$	$\hat{\beta}^{SR}_{CCE}$	$\hat{\kappa}$	$\hat{\mu}$
Whole Sample	*0.80*	*-4%*	*0.80*	*-4%*	*0.82*	*-1%*	0.83 (0.03)	*0.65*	*-10%*	*0.74*	*+2%*	*0.72*	*-1%*	0.73 (0.02)	-0.40	0.30
High + Mid Inc	*0.80*	*-5%*	*0.85*	*+1%*	*0.84*	*0%*	0.84 (0.03)	*0.56*	*-20%*	*0.73*	*+3%*	*0.69*	*-2%*	0.71 (0.03)	-0.38	0.41
High Inc	*0.71*	*-12%*	*0.83*	*+2%*	*0.81*	*0%*	0.81 (0.06)	*0.25*	*-59%*	*0.60*	*0%*	*0.55*	*-9%*	0.60 (0.04)	-0.28	0.91
OECD	*0.75*	*-12%*	*0.88*	*+4%*	*0.86*	*+1%*	0.85 (0.07)	*0.76*	*+16%*	*0.68*	*+5%*	*0.65*	*0%*	0.65 (0.04)	-0.24	0.96
Developed	*0.67*	*–16%*	*0.86*	*+7%*	*0.78*	*-3%*	0.80 (0.05)	*0.16*	*-73%*	*0.57*	*-5%*	*0.53*	*-13%*	0.61 (0.04)	-0.27	0.88

Note: Country group definitions follow those used by the World Bank and the United Nations (developed countries have a Human Development Index value in the top quartile of the HDI distribution). The notes in Table 2 list the model definitions. In the DEM, HYB, and CCE models the individual $\hat{\beta}_i$ are aggregated using equation (15). Standard errors in parentheses. Full international risk sharing implies H_0: $\beta = 0$. All β estimates are different from 0 at the 5% level of marginal significance, but only inference based on the CCE methodology is valid. $\pm$ indicates the percentage difference between the estimate in the preceding column and the respective $\hat{\beta}^{LR}_{CCE}$ or $\hat{\beta}^{SR}_{CCE}$ estimate. $\hat{\kappa}$ denotes the estimated speed-of-adjustment coefficient in the error-correction model. $\hat{\mu}$ denotes the mean adjustment lag computed as $\hat{\mu} = (1 - \frac{\hat{\beta}^{SR}_{CCE}}{\hat{\beta}^{LR}_{CCE}})/(-\hat{\kappa})$ based on Fuleky and Ventura (2016).

displays the estimates of risk sharing behavior for the whole sample and several of its subsets categorized by income and development[3]. In line with earlier studies (such as Becker and Hoffmann, 2006), our CCE results indicate that consumption tends to be affected by idiosyncratic shocks in both the long and the short run, and the extent of risk sharing tends to be higher in the short run. The fraction of idiosyncratic variation in GDP channelled to consumption is slightly above 0.80 in the long run, while it ranges between 0.60-0.73 in the short-run.

Risk sharing in the long run does not exhibit statistically significant variation across subsets of countries. However, our short run CCE results reveal a geo-economic pattern that is similar to the one found by Kose et al. (2009) who analyzed 69 developing and developed countries over the 1960-2004 period. In particular, falling to 0.61 for developed countries, the $\hat{\beta}^{SR}_{CCE}$ estimates are inversely related to the level of development. This signals a greater capacity of developed economies to insure against idiosyncratic risk as they tend to have better access to well functioning credit and capital markets. Moreover, although income level is not necessarily a good approximation to the degree of openness, the $\hat{\beta}^{SR}_{CCE}$ estimates support the notion that higher income countries tend to enjoy a greater degree of risk sharing in the short run.

The conventional estimates are affected by various biases, and as the $\pm$ columns in Table 4 indicate, the gap between the fixed effects estimates and the CCE estimates can be substantial. The fixed effects estimator in (4) will produce different results from the mean-group estimator in (3) if $\hat{\beta}_{DEM,i}$ is correlated with the variance of demeaned income, $S_i = Var(y_{it} - \bar{y}_t)$ (see Coakley et al., 2001). We find that the correlation $Cor(\hat{\beta}^{LR}_{DEM}, S)$ across i is negative, and consequently $\hat{\beta}^{LR}_{FE}$ tends to be lower than $\hat{\beta}^{LR}_{DEM}$. The negative correlation between $\hat{\beta}$ and S implies that countries with high income variation tend to smooth their consumption more than countries with low

[3]The estimation results for each individual country are available in the Appendix.

income variation, which is consistent with the conclusion put forth by Browning and Collado (2001). In contrast, the β^{LR}_{DEM} estimates tend to exhibit an upward bias due to lingering global shocks shared by $c_{it} - \bar{c}_t$ and $y_{it} - \bar{y}_t$.

The β^{LR}_{HYB} estimates tend to be very close to the β^{LR}_{CCE} ones due to the high correlation of aggregate consumption and income, or $\bar{c}_t$ and $\bar{y}_t$, respectively. Having at least one of these aggregate measures as control variables in the model alleviates some of the problems associated with the DEM model. Specifically, the HYB residuals do not contain unit roots. However, the rejection of the cross-sectional independence test in Table 2 indicates that $\bar{c}_t$ alone is not able to fully eliminate common shocks and isolate idiosyncratic fluctuations in equation (7). And, although the mean-group DEM and HYB estimates are statistically indistinguishable from the CCE ones in Table 4, the results for individual countries listed in the Appendix exhibit somewhat greater divergence.

Our mean-group CCE estimates for OECD countries, $\hat{\beta}^{LR}_{CCE} = 0.85$ and $\hat{\beta}^{SR}_{CCE} = 0.65$, fall somewhat below the respective conventional estimates of about 0.9 and 0.7 obtained by Leibrecht and Scharler (2008), who—albeit relying on homogeneity assumptions—also used an error correction model. However, our estimated speed of equilibrium-error correction, $\hat{\kappa} = -0.24$, deviates from their -0.1 estimate by a larger margin. Consequently, the mean adjustment lag (computed as $\hat{\mu} = (1 - \frac{\hat{\beta}^{SR}_{CCE}}{\hat{\beta}^{LR}_{CCE}})/(-\hat{\kappa})$ based on Fuleky and Ventura, 2016) indicates that in OECD countries an idiosyncratic income shock exerts its effect on consumption within about a year according to our study as opposed to about two years according to the results of Leibrecht and Scharler (2008). The last column of Table 4 illustrates the direct relationship between the mean adjustment lag and the level of development. It is the longest in OECD countries, where consumption appears to react to idiosyncratic income shocks more slowly, perhaps due to a robust institutional framework, abundance of consumption

Table 5: CCE Coefficient Estimates for Subsamples, in Subperiods

	$\hat{\beta}^{LR}_{CCE}$		$\hat{\beta}^{SR}_{CCE}$	
Country Group	1970-1989	1990-2014	1970-1989	1990-2014
Whole Sample	0.80 (0.04)	0.77 (0.04)	0.71 (0.04)	0.71 (0.04)
High + Mid Inc	0.80 (0.04)	0.78 (0.04)	0.67 (0.05)	0.68 (0.04)
High Inc	0.78 (0.05)	0.63 (0.07)	0.56 (0.05)	0.51 (0.06)
OECD	0.74 (0.05)	0.76 (0.06)	0.58 (0.04)	0.62 (0.06)
Developed	0.73 (0.06)	0.68 (0.07)	0.55 (0.05)	0.56 (0.06)

Note: See also the notes in Table 4. The hypothesis that the extent of risk sharing remained the same in the 1990-2014 subperiod as in the 1970-1989 one ($H_0 : \hat{\beta}_{1970-1989} = \hat{\beta}_{1990-2014},\ H_1 : \hat{\beta}_{1970-1989} \neq \hat{\beta}_{1990-2014}$) cannot be rejected for any country group at the 5% level of marginal significance.

smoothing opportunities, or direct access to financial markets.

5.3 Risk Sharing in the Globalization Era

Table 5 allows us to contribute to the debate on whether increased financial globalization, an impressive surge in flows of real and financial assets across countries, has brought about more, or better, insurance opportunities. Economic theory does not necessarily imply that this should be the case. Whether asset trading fosters risk sharing, crucially depends on the co-movements of domestic and foreign asset returns. The benefit will be limited if, due to geographic, political or cultural proximity, countries only engage in asset trading with partners that are affected by similar shocks. Neither will procyclical foreign credit—abundant in booms and scarce in busts—contribute to international risk sharing. On the other hand, asset trading between countries that experience asymmetric shocks is expected to result in a greater degree of insurance.

Empirical evidence in favor of a diversification motive in asset trading is mixed.

For example, Lane and Milesi-Ferretti (2008) found little evidence that gains from diversification drive bilateral cross-country asset holdings. Instead, they observed that investors tend to hold equity in destinations with similar business cycle and stock market behavior. On the other hand, Pericoli et al. (2013), using the same data but resorting to a (panel) fractional regression model for investment shares, concluded that asset trading does appear to be affected by an incentive to diversify risk. Also, Sorensen et al. (2007) documented that during the 1990s a decline in home bias was associated with an increase in risk sharing in OECD countries. Kose et al. (2009) found that, while industrial countries have attained higher levels of risk sharing during the recent period of globalization, developing countries have been mostly shut out of these benefits. They attributed this result to the composition of capital flows, with external debt preventing many emerging economies to efficiently share risks. Similarly to Bai and Zhang (2012), they suggested that the dichotomy can be explained by the existence of threshold mechanisms, whereby only countries reaching a certain level of financial development reap the benefits of financial globalization.

To get some insight as to whether the level of risk sharing has changed over time, we have repeated our analysis for two subperiods, one running from 1970 to 1989, the other covering the period 1990 to 2014. Because the reduction in sample size affects the statistical properties of the two subperiods about equally, the estimates corresponding to each subsample can be compared to each other, revealing some interesting results.[4] Specifically, none of the country groups experienced a significant improvement in international consumption risk sharing in the financial globalization era, whether in the long or in the short run.

These conclusions stand in stark contrast with those of Artis and Hoffmann (2012), who estimated the fixed effects regression (4) using a subset of OECD countries that

[4]Our results do not materially change if we eliminate the impact of the Great Recession by limiting the time horizon to 2007.

excluded Chile, Hungary, Israel, Korea, Mexico, Poland, and Turkey, and found a significant increase in long-run risk sharing. Using the dataset and estimator in their study, we replicated the results of Artis and Hoffmann (2012), which were $\hat{\beta}^{LR}_{FE} = 0.98$ for 1960-1990 and $\hat{\beta}^{LR}_{FE} = 0.63$ for 1990-2004. However, once we allowed for general heterogeneity in the model, their conclusions suggesting an increase in risk sharing broke down: we obtained $\hat{\beta}^{LR}_{CCE} = 0.93$ for 1960-1990 and $\hat{\beta}^{LR}_{CCE} = 0.94$ for 1990-2004. Our results appear to support the view that while there has been an increase in the volume of financial transactions across the world, as of now, globalization has not triggered an increase in international consumption risk sharing.

6 Conclusion

We study the impact of cross-sectional heterogeneity on conventional tests of international risk sharing. Relying on the restrictive assumption of symmetric country characteristics, the existing literature typically employs cross-sectional demeaning to filter out global shocks from the consumption and income panels. We find that that approach is not able to eliminate common factors from heterogeneous data sets, and consequently the coefficient estimate is affected by the correlation between aggregate consumption and aggregate income. Moreover, imposing pooled estimation distorts the coefficient estimates due to a correlation between the extent of consumption smoothing and the variation in income.

Inspired by the inadequacy of the conventional approach to isolate idiosyncratic fluctuations in a diverse set of 120 countries, we propose an alternative approach. We control for global factors via heterogeneous loading coefficients within an unobserved components framework that parallels the CCE methodology of Pesaran (2006) and Kapetanios et al. (2011). While the mean-group estimates by the demeaning (DEM)

and hybrid (HYB) approaches are statistically indistinguishable from the CCE estimates, the fixed effects (FE) estimates based on the pooled data set differ from the proposed method by up to 73%. This implies that inappropriately imposing a uniform β parameter across countries, or ignoring bullet point 3 in Section 1, is the most critical mistake. But if the objective is to analyze the relationship between idiosyncratic fluctuations in income and consumption, then any of the cross-sectional demeaning approaches that ignore bullet points 1 and 2 in the Introduction will fail. Once the focus shits from mean-group aggregates to results for individual countries, the difference between conventional estimates and CCE ones becomes more pronounced.

Our results confirm the lack of evidence for full risk sharing, with the degree of risk sharing being lower in the long run. We show that developed economies, benefitting from more opportunities to insure against risk, are affected by idiosyncratic shocks more slowly and to a lesser extent. Finally, in contrast to some earlier empirical findings, we do not detect any evidence of a recent increase in international risk sharing once we appropriately control for cross-sectional heterogeneity in the data.

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Appendix
(to be made available online)

Table A1: Coefficient Estimates for Individual Countries

Country	$\hat{\beta}^{LR}_{DEM}$	$\hat{\beta}^{LR}_{HYB}$	$\hat{\beta}^{LR}_{CCE}$	$\hat{\beta}^{SR}_{DEM}$	$\hat{\beta}^{SR}_{HYB}$	$\hat{\beta}^{SR}_{CCE}$	Country	$\hat{\beta}^{LR}_{DEM}$	$\hat{\beta}^{LR}_{HYB}$	$\hat{\beta}^{LR}_{CCE}$	$\hat{\beta}^{SR}_{DEM}$	$\hat{\beta}^{SR}_{HYB}$	$\hat{\beta}^{SR}_{CCE}$
AGO	1.40	1.43	1.46	0.92	1.34	1.34	EGY	0.70	0.79	0.81	0.59	0.34	0.37
ALB	0.45	0.07	0.07	0.29	0.26	0.28	ESP	1.10	1.07	1.06	1.03	0.96	0.92
ARG	0.71	0.94	0.93	1.15	1.12	1.11	ETH	0.07	0.33	0.34	0.36	0.31	0.31
AUS	1.03	0.94	0.92	0.50	0.36	0.37	FIN	0.85	0.89	0.89	0.68	0.61	0.60
AUT	0.89	1.02	1.02	0.89	0.81	0.84	FRA	1.00	1.00	1.00	0.79	0.73	0.75
BDI	0.78	0.88	0.98	0.84	0.88	0.94	GAB	0.72	0.59	0.37	0.24	0.24	0.28
BEL	0.85	0.98	0.97	0.78	0.72	0.73	GBR	1.19	1.06	1.03	0.91	0.81	0.77
BEN	1.50	-0.19	-0.15	0.36	0.28	0.29	GHA	0.97	0.92	0.80	1.17	1.21	1.13
BFA	0.73	1.28	1.49	1.24	1.18	1.17	GIN	0.87	0.39	0.41	0.57	0.75	0.77
BGD	0.16	1.33	1.31	0.58	0.62	0.62	GMB	1.24	1.60	1.70	0.77	0.95	1.20
BGR	0.88	1.15	1.12	0.84	0.93	0.96	GNB	0.71	0.76	0.75	1.02	1.01	1.04
BOL	0.90	0.73	0.69	0.58	0.71	0.78	GRC	0.85	1.12	1.13	0.74	0.79	0.79
BRA	0.60	0.61	0.81	0.80	0.79	0.75	GTM	0.74	0.73	0.72	0.79	0.81	0.77
BWA	0.80	0.64	0.67	0.53	0.55	0.53	HKG	1.02	1.13	1.14	0.79	0.72	0.75
CAF	0.90	0.72	0.71	0.86	0.87	0.87	HND	0.89	0.95	1.01	0.66	0.91	0.84
CAN	0.53	0.52	0.53	0.55	0.50	0.47	HTI	0.29	0.27	0.57	0.65	0.61	0.55
CHE	1.19	0.87	0.87	0.57	0.47	0.45	HUN	0.71	0.79	0.83	0.72	0.82	0.79
CHL	0.93	0.93	0.74	0.99	0.91	0.95	IDN	1.18	1.11	1.13	0.79	0.67	0.68
CHN	1.35	1.51	1.46	0.52	0.49	0.53	IND	0.76	0.80	0.79	0.67	0.66	0.71
CIV	0.95	1.04	1.07	1.15	1.19	1.23	IRL	0.52	0.70	0.69	0.76	0.68	0.66
CMR	0.87	0.85	0.90	0.63	0.66	0.70	IRN	0.24	0.33	0.45	0.44	0.33	0.30
COD	0.80	0.86	0.94	1.08	1.37	1.18	IRQ	-0.12	0.06	0.07	0.28	0.28	0.44
COG	0.94	0.93	0.72	0.60	0.56	0.54	ISR	0.71	0.93	0.91	0.80	0.47	0.44
COL	0.71	0.81	0.91	0.85	0.70	0.70	ITA	0.98	0.98	0.99	0.90	0.80	0.81
CRI	-0.07	1.20	1.21	1.00	1.04	1.02	JAM	0.57	1.02	0.76	0.71	0.56	0.58
DEU	0.90	0.98	0.99	0.55	0.43	0.45	JOR	1.23	1.12	1.23	0.86	0.90	0.94
DNK	0.68	0.68	0.68	0.82	0.78	0.71	JPN	0.88	0.85	0.88	0.73	0.61	0.63
DOM	0.94	0.70	0.64	1.11	1.16	1.15	KEN	0.68	0.09	0.85	1.11	1.07	1.04
DZA	1.32	1.51	1.18	0.33	0.44	0.45	KHM	0.76	0.93	1.00	0.90	0.87	0.88
ECU	0.93	0.86	0.71	0.33	0.37	0.40	KOR	0.73	0.78	0.79	0.81	0.71	0.73

Note: See also the notes in Table 4.

Table A1 continued: Coefficient Estimates for Individual Countries

Country	$\hat{\beta}^{LR}_{DEM}$	$\hat{\beta}^{LR}_{HYB}$	$\hat{\beta}^{LR}_{CCE}$	$\hat{\beta}^{SR}_{DEM}$	$\hat{\beta}^{SR}_{HYB}$	$\hat{\beta}^{SR}_{CCE}$	Country	$\hat{\beta}^{LR}_{DEM}$	$\hat{\beta}^{LR}_{HYB}$	$\hat{\beta}^{LR}_{CCE}$	$\hat{\beta}^{SR}_{DEM}$	$\hat{\beta}^{SR}_{HYB}$	$\hat{\beta}^{SR}_{CCE}$
KWT	-0.02	-0.16	-0.18	-0.40	-0.46	-0.30	POL	1.03	0.90	0.62	0.80	0.83	0.84
LAO	0.64	1.05	1.06	0.76	0.76	0.80	PRT	0.96	0.93	0.92	0.82	0.76	0.70
LBN	0.87	0.74	0.86	0.93	0.90	0.89	PRY	0.78	0.77	0.97	0.39	0.30	0.31
LBR	0.87	1.05	1.04	1.03	1.13	1.06	PSE	0.67	0.69	0.71	0.62	0.60	0.65
LKA	1.18	1.26	1.26	0.81	0.91	0.93	ROU	0.90	0.30	0.44	0.66	0.60	0.54
LSO	0.62	0.93	0.90	0.42	0.56	0.55	RWA	0.99	0.88	0.83	0.51	0.56	0.54
MAR	1.01	1.04	1.28	0.76	0.82	0.89	SAU	1.11	1.28	1.28	0.77	0.80	0.75
MDG	1.12	1.01	0.99	0.63	0.66	0.65	SDN	0.51	0.57	0.85	0.81	0.74	0.83
MEX	0.72	0.65	0.74	1.00	0.98	0.99	SEN	1.06	0.90	0.94	0.42	0.49	0.51
MLI	0.45	0.80	0.78	0.26	0.40	0.39	SGP	0.66	0.75	0.79	0.59	0.53	0.55
MMR	0.63	0.96	1.05	0.97	1.00	0.93	SLE	0.89	0.78	0.79	0.81	1.01	0.99
MNG	1.31	0.94	1.01	1.33	0.64	0.52	SLV	1.10	1.15	1.01	1.07	1.09	1.11
MOZ	0.46	0.80	1.10	0.68	0.60	0.69	SWE	0.39	0.68	0.69	0.53	0.43	0.44
MRT	0.71	0.60	0.63	0.87	0.86	0.86	SYR	0.87	0.70	0.65	0.86	0.83	0.85
MUS	0.95	0.93	0.92	0.93	0.82	0.80	TCD	0.54	0.44	0.48	0.70	0.71	0.78
MWI	0.77	0.32	0.40	0.46	0.57	0.58	TGO	0.63	0.38	0.38	0.72	0.46	0.48
MYS	0.90	0.57	0.57	0.99	1.03	1.03	THA	0.85	0.71	0.72	0.79	0.69	0.73
NAM	0.68	1.06	1.06	0.47	0.29	0.21	TTO	0.49	0.77	0.95	0.73	0.33	0.54
NER	0.83	0.56	0.56	1.17	0.90	0.88	TUN	1.23	1.00	0.98	0.44	0.49	0.51
NGA	1.21	1.21	1.09	0.67	0.62	0.43	TUR	0.72	0.62	0.60	0.80	0.66	0.65
NIC	0.77	0.76	0.70	0.64	0.69	0.64	TWN	0.90	1.12	1.13	0.85	0.66	0.75
NLD	0.75	0.85	0.85	0.88	0.75	0.73	TZA	0.09	0.35	0.18	0.32	0.91	0.88
NOR	0.72	0.58	0.59	0.62	0.56	0.58	UGA	0.97	0.99	0.85	0.99	0.99	1.05
NPL	1.05	1.08	0.99	0.97	0.89	0.93	URY	1.19	1.20	1.27	1.16	1.09	1.08
NZL	0.75	0.60	0.39	0.68	0.63	0.47	USA	0.88	0.89	0.88	0.85	0.70	0.68
OMN	0.64	0.52	0.43	0.85	0.86	0.81	VEN	0.50	0.93	0.83	0.90	1.03	0.94
PAK	0.66	0.70	0.71	0.71	0.70	0.95	VNM	0.67	0.59	0.62	0.83	0.84	0.88
PAN	0.24	0.32	0.37	0.47	0.45	0.52	ZAF	0.42	0.46	0.46	0.89	0.89	0.88
PER	1.18	1.03	1.02	1.05	1.03	1.03	ZMB	1.45	1.16	1.22	0.59	0.60	0.70
PHL	0.31	0.29	0.37	0.39	0.45	0.45	ZWE	0.47	1.21	1.15	1.44	1.59	1.44

Note: See also the notes in Table 4.