ELSEVIER European Economic Review 41 (1997) 1079–1109

EUROPEAN ECONOMIC REVIEW

Are countries with official international restrictions 'liquidity constrained'?

Karen K. Lewis *

University of Pennsylvania, Wharton School, Department of Finance, 3620 Locust Walk, Philadelphia, PA 19104-6367, USA
NBER, Cambridge, MA 02138, USA

Received 15 June 1996; revised 15 November 1996

Abstract

In this paper, I empirically examine consumption smoothing behavior across a broad group of countries using a unique data set that indicates whether residents in a country face an official government restriction. I then ask whether the ex ante consumption movements among restricted countries differ from those of unrestricted countries. To gauge the departure from standard consumption smoothing, I use the Campbell and Mankiw ('Consumption, income, and interest rates: Reinterpreting the time series evidence', In: O.J. Blanchard and S. Fischer, eds., NBER macroeconomics annual, 1989 (MIT Press, Cambridge, MA, 1989) and 'The response of consumption to income: A cross-country investigation', European Economic Review 35, 723–756, 1991) approach of regressing consumption growth on income growth and instrumenting with lagged variables. Interestingly, I find that consumption growth for residents in countries that impose international restrictions have a significantly higher coefficient on income growth than do residents in countries without those restrictions. Thus, a greater proportion of consumers facing international restrictions appear to act as though they are liquidity constrained according to the Campbell and Mankiw approach. I also discuss alternative interpretations that do not depend upon liquidity constraints. © 1997 Elsevier Science B.V.

JEL classification: F3

Keywords: Consumption behavior; Output growth; International restrictions

* Tel.: (+1) 215-898.7637; e-mail: lewisk@wharton.upenn.edu.

PII S0014-2921(97)00056-1

1. Introduction

Consumption is clearly an important economic variable as its prominence in literatures spanning macroeconomics, asset pricing, and international economics testifies. In this paper, I empirically examine intertemporal consumption smoothing behavior across a broad group of countries to address some recent issues posed by the intersection of these three literatures. For this purpose, I employ a unique data set that indicates whether residents in a country face an official government restriction. I then ask whether the ex ante consumption movements among restricted countries differ from those of unrestricted countries.

To address this question, I modify the Zeldes (1989) approach of comparing the consumption Euler equations of individuals who are unlikely to be liquidity constrained with those of other individuals. I need to modify the approach because, while Zeldes found that Euler equations were not rejected for the unconstrained individuals, it is well known that Euler equations are rejected using aggregate data. [1] To gauge the degree of liquidity constraints, therefore, I use the practice made popular by Campbell and Mankiw (1989, 1991) (hereafter, CM) of regressing consumption growth on income growth and instrumenting with lagged variables. CM interpret the coefficient on income growth as the proportion of individuals who are liquidity constrained. I begin with this interpretation despite some important caveats. I later provide alternative interpretations that do not depend upon liquidity constraints.

The empirical results are striking. Generally, I find that consumption growth for residents in countries that impose international restrictions have a significantly higher coefficient on income growth than do residents in countries without those restrictions. Thus, a greater proportion of consumers facing international restrictions appear to act as though they are liquidity constrained according to the CM interpretation.

There are at least two reasons why these results are interesting. The primary reason and motivation for this paper comes from recent research in both international finance and business cycle research finding that countries do not optimally share risk. The international finance literature has documented the well-known 'home bias' puzzle, the phenomenon that domestic residents strongly bias their equity holdings toward domestic stocks foregoing the lower variance and potentially higher returns from a portfolio with more foreign stocks. [2] The international business cycle literature has found a dual puzzle by examining consumption behavior across countries. [3] One explanation proposed for this behavior is that

[1] See Deaton (1992), for example.

[2] See for example, French and Poterba (1991), Baxter and Jermann (1995), and the discussion in Lewis (1995).

[3] In contrast to the standard predictions of complete markets, Backus et al. (1992) and others show that consumption growth rates have a lower correlation across countries than output correlations.

markets are incomplete. [4] Indeed, Lewis (1996) finds empirical evidence that international capital market restrictions affect the equilibrium degree of consumption risk-sharing.

This observation in the international market appears to be somewhat at odds with the prevalent view in the asset pricing literature, however. For example, Heaton and Lucas (1992, 1995) and Telmer (1993) show that individuals facing frictions such as transactions costs in one capital market can effectively get around them by transacting in another market. Thus, applying the idea to the international context, one would expect that restrictions in foreign equity acquisitions would have little effect upon equilibrium consumption sharing if international investors have access to international borrowing and lending.

The evidence in this paper may therefore shed some light on this apparent inconsistency. Using the CM interpretation, these results suggest that countries facing restrictions in acquiring foreign stocks are also unable to borrow and lend internationally on the same terms as others. [5] Thus, residents in these countries appear to face frictions in many capital markets at once and, therefore, may be unable to duplicate consumption insurance through other markets.

The second main reason why these results are interesting is that they document new evidence on market restrictions and consumption behavior in a growing literature on this subject. While the results in this paper represent the first attempt to statistically test for intertemporal consumption differences across countries, several other studies have examined the informal relationship between capital market frictions and differences in liquidity constraints across countries or over time. Jappelli and Pagano (1989) examine the sensitivity of consumption to income for seven countries. They use credit market variables such as consumer debt and mortgages to argue that the differences in income sensitivity across countries arise from differences in liquidity constraints. In a similar vein, Bayoumi (1993) examines the consumption sensitivity to income in the UK over time and considers whether this time-variation can be attributed to variation in financial market regulations. Relative to this literature, I use a new set of international restrictions measures and introduce a new empirical approach that allows direct testing for differences in 'liquidity constraints' across countries.

The plan of the paper is as follows. Section 2 develops the tests and data to be used in the study. Section 3 describes the results. Section 4 provides other possible explanations for the results. Concluding remarks follow.

[4] For example, Baxter and Crucini (1995) examine an open economy model in which investors in each countries can only borrow and lend, but cannot acquire shares in each other's output. Explanations for this puzzle also focus upon non-tradeable goods and leisure, a possibility I consider in Section 5. See Stockman and Tesar (1995), Tesar (1993), Devereux et al. (1992), and Baxter et al. (1995).

[5] This inability to borrow or lend may also arise from domestic credit restrictions. Without a data set on domestic credit restrictions on the broad set of countries examined here, I am unable to identify the international from domestic credit restrictions.

2. Testing for differences in 'liquidity constraints' across internationally restricted countries

A large literature has tested the implications of the consumption Euler equation. Furthermore, a number of studies beginning with Zeldes (1989) have tested for differences in the implications of the Euler equations among different groups. In this paper, I follow this tradition by examining the Euler equation differences across groups that are more likely to face capital market restrictions.

This paper differs from the standard literature in several ways, however. First, I test for differences across countries. Second, I employ a unique panel data set of six official restrictions that may potentially proxy for market restrictions. While the data may not even proxy accurately for these restrictions, they allow an original comparison across a broad set of countries and permit the data to speak about whether the restrictions are significant. Third, I develop a new method of testing for differences in Euler equations across countries. Specifically, I interact the capital market restrictions measures with the Euler equations and test for statistically significant differences across countries. Below, I develop these tests and describe the data.

2.1. Intertemporal substitution and the 'bonds-only' assumption in international macroeconomics

Consider the decision of individuals in a given country of whether to consume today or invest in an asset with a possibly risky return. A standard approach in the international macroeconomics literature is to assume that each country is populated with a representative agent, an assumption I follow below. Then each country's agent maximizes utility functions $u(C_t^j)$ where j indexes the countries, $j = 1, \ldots, J$ and C_t^j is an aggregate consumption good at time t for this country j. Clearly, this problem is completely analogous to the optimal intertemporal consumption decision that has been extensively studied in asset pricing and macroeconomics.[6] Defining the return on any asset as R_t, this decision implies the standard Euler equation,

$$\mathrm{E}_t\left[R_t \rho^j u'\left(C_{t+1}^j\right) / u'\left(C_t^j\right)\right] = 1, \tag{1}$$

where ρ^j is the discount rate for country j.

The intertemporal approach to the current account assumes further that countries have access to a risk-free international real bond market for borrowing and lending.[7] I define the rate on this risk-free bond as R_t^{w}. Assuming consumption

[6] For surveys using consumption, see Hall (1990) or Deaton (1992). For an early study exploiting this relationship in empirical asset pricing, see Hansen and Hodrick (1980).

[7] See for example Baxter and Crucini (1995) or Obstfeld and Rogoff (1995).

growth is log-normally distributed and conditionally homoscedastic, that utility is isoelastic, $u(C_t^j) = (C_t^j)^{(1-\gamma)}/(1-\gamma)$, and substituting actual for expected consumption growth, the Euler equation can be rewritten as [8]

$$\Delta c_{t+1}^j = \left(\tfrac{1}{2}\sigma_j^2 + (1/\gamma)\ln(\rho^j)\right) + (1/\gamma) r_t^{\mathrm{w}} + \epsilon_{t+1}^j. \tag{2}$$

where lower case letters represent the logarithm of the variables and $\epsilon_{t+1}^j \equiv \Delta c_{t+1}^j - \mathrm{E}_t \Delta c_{t+1}^j$.

Euler equations of the form given in Eq. (2) have been estimated and tested in standard forms such as the regression

$$\Delta c_{t+1}^j = \theta_0^j + \theta_1 r_t^w + \beta X_t^j + \epsilon_{t+1}^j, \tag{3}$$

where $\theta_0^j = (\frac{1}{2}\gamma\sigma_j^2 + (1/\gamma)\ln(\rho^j))$, a fixed country effect and $\theta_1 = (1/\gamma)$. X_t^j is defined as any variable in the time t information set so that under the null hypothesis of intertemporal substitution, $\beta = 0$. Campbell and Mankiw (1989, 1991) treat X_t^j as the logarithm of actual income growth, Δy_{t+1}^j, instrumented with information in the time t information set. They interpret the coefficient β as the share of consumers who are liquidity constrained.

To recall this interpretation, I briefly outline the Campbell–Mankiw story here and refer readers to the original papers for further details. There are two groups of individuals in the economy. The first group is not liquidity constrained and therefore follow Eq. (2). The second group of individuals cannot borrow or lend and are so-called 'rule-of-thumb' consumers. These individuals consume all of their income so that instead of Eq. (2), their consumption follows the process

$$\Delta c_{t+1}^{\mathrm{RoT},j} = \Delta y_{t+1}^j + \epsilon_{t+1}^j, \tag{4}$$

where the superscript 'RoT' indexes 'rule-of-thumb' consumers and y_t^j is the logarithm of income for j at time t. Next, the story assumes that the proportion of 'rule-of-thumb' consumers in the economy is constant and equal to β. Therefore, summing across individuals in the economy to obtain aggregate consumption implies a share-weighted average of Eq. (2) and Eq. (4):

$$\begin{aligned}\Delta c_{t+1}^j &= (1-\beta)\left[\tfrac{1}{2}\sigma_j^2 + (1/\gamma)\ln(\rho^j)) + (1/\gamma) r_t^{\mathrm{w}}\right] + \beta \Delta y_{t+1}^j + \epsilon_{t+1}^j \\ &= \theta_0^j + \theta_1 r_t^{\mathrm{w}} + \beta X_t^j + \epsilon_{t+1}^j, \end{aligned} \tag{5}$$

which has the same form as Eq. (3) but where now β is explicitly the proportion of individuals who are 'rule-of-thumb' consumers.

[8] Below, I relax the assumption that consumption is conditionally homoscedastic. The basic results for the sensitivity of consumption to income are robust to relaxing this assumption.

There are at least two problems with this interpretation, however. [9] First, there is no reason to expect that liquidity constrained consumers would consume all of their income according to the 'rule-of-thumb' as in Eq. (4). Indeed, as Zeldes (1989) has argued, liquidity constrained individuals may have a greater precautionary savings motive during high income periods, since they realize that they will be unable to borrow during low income periods. Therefore, they would not consume all of their income.

Second, although this framework assumes that the proportion β of individuals who are 'rule-of-thumb' consumers are constant, this proportion would necessarily vary over time if these individuals are always in the 'rule-of-thumb' group. Put differently, summing over the growth rate of individuals as done to obtain Eq. (5) is not the same as the growth rate of the aggregate consumption; i.e., for individual i consumption, C^i, within an economy, $\Delta \ln(\Sigma_i C^i) \neq \Sigma_i \Delta \ln(C^i)$.

Despite these problems with interpretation, I use the Campbell–Mankiw regression approach because it provides a useful summary of the relationships in the data. In particular, these regressions show the responsiveness of consumption growth to predictable income growth. Even if the CM interpretation is not valid, the finding that these responses differ between restricted and unrestricted countries is an interesting empirical result and may have alternative interpretations. As I discuss in Section 4, these alternative explanations include differences in measurement error, habit persistence, non-separabilities between tradeables and non-tradeables, and time-variation in the consumption variances, σ_j^2.

2.2. Structure of the tests

Zeldes (1989) shows how rejections of the hypothesis that $\beta = 0$ may provide evidence about liquidity constraints. When individuals are liquidity constrained, then the left-hand side of Eq. (5) will have an additional variable that reflects the shadow value of the restriction. He provides a test that segments the sample into two groups based upon whether they are likely to be constrained. [10]

[9] A third problem is that if consumers follow a different 'rule-of-thumb' than the assumed one, it is not possible to identify the proportion who follow this rule. For example, if instead of consuming all of income as in Eq. (4) they consume a proportion b of their income, then summing across the group of unconstrained and 'rule-of-thumb' consumers would imply that Eq. (5) is

$$\Delta c_{t+1}^j = \theta_0^j + \theta_1 r_t^{\mathrm{w}} + \beta b \Delta y_{t+1}^j + \epsilon_{t+1}^j,$$

where now the coefficient on income is the product of the proportion of individuals who are 'rule-of-thumb' and the 'rule-of-thumb' parameter itself, b. Thus, the estimate on income will be a product of these two effects that could not be identified separately.

[10] Zeldes used household data splitting the sample based upon the wealth of the household. Jappelli et al. (1995) use more direct information on borrowing constraints using two different data sets to model the split endogenously.

In the international context, I consider a decomposition based upon whether residents of individual countries face governmental restrictions for international trade during the period or not. For this purpose, I define

$$D(j,t) \quad = 1 \quad \text{if country } j \text{ is restricted at time } t$$
$$\qquad\quad = 0 \quad \text{if country } j \text{ is not restricted at time } t.$$

I then consider estimates of the form

$$\Delta c_{t+1}^{j} = \theta_0^{j} + \theta_0(t) + \beta^{\mathrm{R}} D(j,t)\,\Delta y_{t+1}^{j} + \beta^{\mathrm{U}}(1 - D(j,t))\,\Delta y_{t+1}^{j} + \epsilon_{t+1}^{j}, \tag{6}$$

where $\theta_0(t)$ is a common fixed time effect across countries at time t and thus captures all variables fixed at time t, including $\theta_1 r_t^{\mathrm{w}}$. Also, the superscripts R refer to restricted and U refer to unrestricted. Under the CM interpretation of β, a restricted country may be expected to have a higher share of liquidity constrained consumers: $\beta^{\mathrm{R}} > \beta^{\mathrm{U}}$. Interestingly, I find this basic pattern in the results below.

2.3. The data

An ideal data set to examine the effects of capital market restrictions across countries would provide direct measures of the effective degree of credit controls across countries. Even such an ideal data set would share the problem faced by other studies in the consumption literature that these measures are likely to be endogenous. [11]

Unfortunately, the available data for market restrictions measures across a broad set of countries are much more crude. In this study, I examine the variables from the International Monetary Fund's *Exchange Restrictions and Exchange Arrangements*. This annual report summarizes the international market restrictions imposed by each country and provides a series of dummy variables for whether the country had a particular restriction in the year. The classifications of these restrictions are in some cases rather unspecific and need not have anything to do with credit market conditions. However, as the literature on financial repression suggests, countries with international restrictions are also likely to have regulations in internal credit markets. [12]

[11] For example, while studies such as Jappelli and Pagano (1989) and Ludvigson (1995) examine the debt levels to determine potential credit tightness, it is not clear whether high debt indicates an economy or household with a higher or lower likelihood of liquidity constraints. That is, high debt may mean a country is nearer a hard constraint, but it may also mean that the country is particularly credit-worthy.

[12] Financial repression occurs when governments restrict international capital markets to provide a larger domestic tax base and greater latitude for domestic credit market regulations. See, for instance, Giovannini and De Melo (1993). Also, for recent descriptions of the effects of liberalization or stabilization policies on domestic internal markets, see Folkerts-Landau and Ito (1995) and Rebelo and Vegh (1995).

Table 1
Average restrictions over the sample period by world group

Group	(1) Capital market	(2) Current account	(3) Export taxes	(4) Import taxes	(5) Import deposit requirements	(6) Bilateral payments
Group of Seven	0.392	0.111	0.360	0.005	0.048	0.000
Continents:						
Africa	0.984	0.727	0.949	0.243	0.097	0.097
North & Central Americas	0.596	0.503	0.636	0.500	0.210	0.086
South America	0.641	0.593	0.952	0.681	0.441	0.400
Asia	0.736	0.551	0.751	0.499	0.336	0.412
Europe	0.752	0.393	0.648	0.205	0.116	0.349
Oceania	0.667	0.185	0.667	0.037	0.037	0.000

Of course, whether these various restrictions affect the ability of individuals to intertemporally consumption smooth is unknown a priori. For this reason, I retain an agnostic approach. I use all of the available restrictions measures and simply ask the data whether the measures of restrictions affect the consumption smoothing behavior.

Thus, I examine the six measures of restrictions provided by the annual report from 1967 to 1992. [13] Table 1 provides the average over the sample period of the number of countries restricted according to these measures broken down by continents and the Group of Seven (G-7). [14] The table lists the measures roughly corresponding to those that affect the largest number of countries moving from left to right.

The first and broadest measure is called 'Restrictions for payments on capital transactions' in the report, referred to here as 'Capital Market' for short. Since this restriction is considered in place when there is any governmental restriction on international capital movements, it is quite general and combines countries with extremely tight capital market restrictions with others that have weaker restrictions. For example, in 1990, both Algeria and Greece were classified as restricted according to this measure. [15] However, the summary description of Algeria indicates that it was extremely difficult to move capital in and out of the country at

[13] The annual report actually contains two additional series. The first series indicates whether a country was in interest arrears over the year. An earlier version of this paper provided evidence using this measure which suggested that countries with interest arrears exhibited significantly different consumption smoothing behavior, consistent with the findings in this paper. However, the data set only begins in 1986 and since the sample period is short, the interest arrears series was omitted in this version. The second series was a 'Bilateral Payments Arrangements' measure that is quite similar to the one used in this paper and was therefore omitted to conserve space.

[14] The G-7 are: the United States, Germany, Japan, the United Kingdom, Italy, France, and Canada.

[15] The data appendix to Lewis (1996) describes these two cases in more detail.

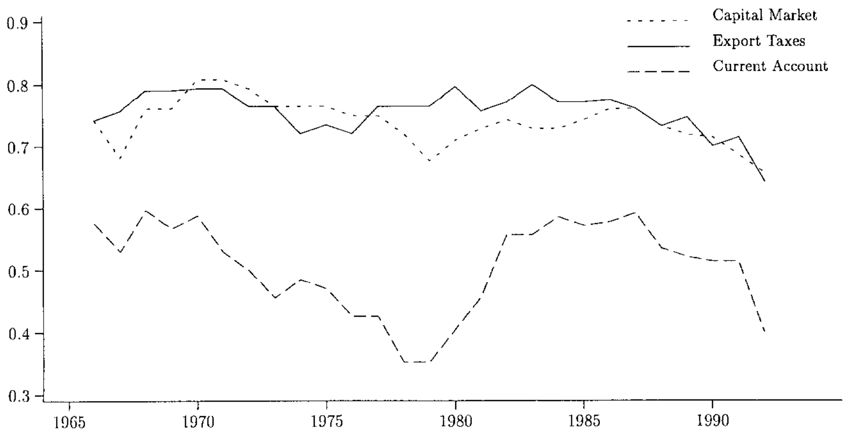

Fig. 1. Proportion of countries with broad-based international restrictions.

all, while for Greece the restriction required domestic firms who wanted to borrow in foreign currency to have debt contracts with maturity of at least six months. Clearly, the cases of these countries represent a rather wide range of capital market restrictions.

The second column reports on 'Restrictions for payments on current account transactions'. As with the 'Capital Market' restriction measure, the 'Current Account' restriction measure covers a large number of restrictions. In this case, however, the restrictions affect the international trade in goods and services as opposed to capital transactions. Table 1 shows that a large proportion of countries across continents are also affected by these measures.

The third column summarizes the continental break-down of 'Export Taxes' by group. This measure indicates whether a country imposes a tax on exporters in the form of requiring export proceeds to be repatriated or surrendered.

The evolution of these first three measures over time are given in Fig. 1. This figure shows that the proportions of all countries facing these types of restrictions are quite high, particularly for the 'Capital Market' and 'Export Taxes' measures. Furthermore, the proportion of countries facing these restrictions have declined only slightly over time. While this fact may appear inconsistent with the conventional wisdom that goods and capital markets have liberalized over time, it is important to recognize that this figure only shows the proportion of countries restricted and does not measure the degree of restriction. Therefore, these measures may miss finer distinctions in the effectiveness of restrictions.

By contrast, the remaining three restrictions are somewhat finer measures that cover fewer of the world's countries. Table 1 shows that the proportion of countries affected by these measures tend to be smaller. Two measures concern import transactions and are given in columns 4 and 5. 'Import Taxes' indicate

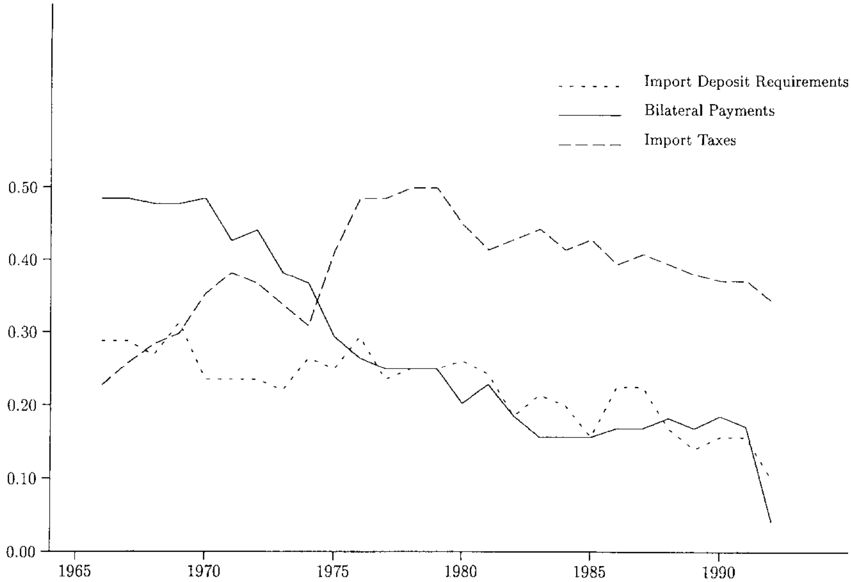

Fig. 2. Proportion of countries with finer-based international restrictions.

countries that require importers to pay surcharges, while 'Advance Import Deposits' imply that importers must post a deposit with the government.

The final measure is called 'Bilateral Payments'. This measure signifies

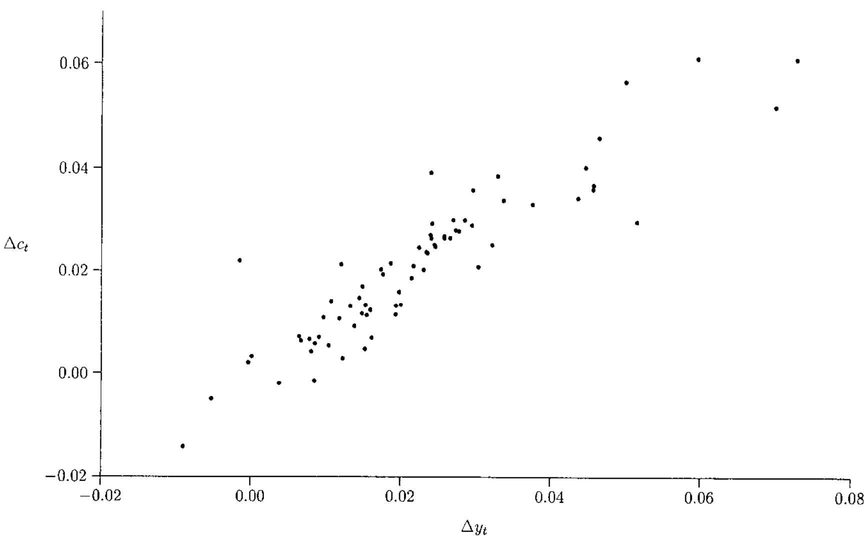

Fig. 3. Plot of average consumption and output growth by country for 1967–1992.

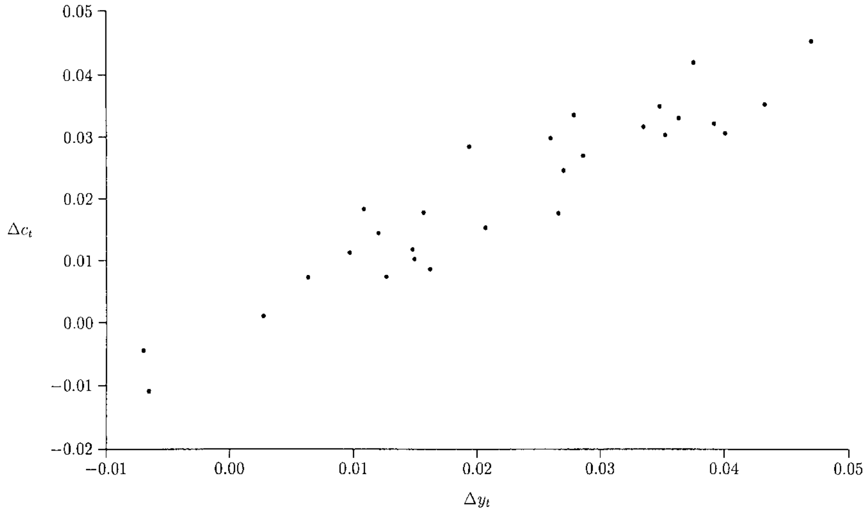

Fig. 4. Plot of average consumption and output growth by year.

whether a country has arrangements for payments with other countries in non-market form such as barter. Since market transactions in internationally traded currencies are typically considered more efficient means of capital movements, bilateral payments arrangements are usually set up as a 'second best' solution. Many of these countries do not have convertible currencies, for instance.

Fig. 2 illustrates how the proportion of countries affected by these three measures have evolved over time. 'Bilateral Payments' show the most marked decline over time, beginning the sample at nearly 50% of the countries and declining to only 4%. The practice of requiring 'Import Deposits' has also declined across countries over time. However, the proportion of countries with some 'Import Taxes' increased in the late seventies before falling somewhat in the 1990s.

For consumption and output data, I follow the international consumption risk sharing literature in using the Summers and Heston (1991) data set. The countries in this group are listed in the data appendix. [16] I use the Penn World Tables, version 5.6 which updates the data through 1992. The consumption and output are in 1985 per capita real terms. [17] Figs. 3 and 4 plot the consumption growth rates

[16] Obstfeld (1994a,b) and Tesar (1995) use these data, for example, dropping countries with data below a given quality, a practice followed here.

[17] In Lewis (1996), I test for the effects of non-separabilities on cross-sectional risk sharing using observations for different groups of countries at five-year intervals. However, intertemporal substitution requires observations for a given country for at least three consecutive periods to provide a current and lagged growth rate of consumption. At most two countries that faced restrictions had three consecutive observations. Clearly, two data points for the restricted group are insufficient for empirical testing.

against the output growth rates averaged, respectively, over time for each country and over country for each year. The figures show a strong positive association between consumption growth and output growth.

For the nominal world interest rate, I use the dollar London interbank offer rate (LIBOR) from the *Bank of England Quarterly Statistics*. This rate is adjusted for price level changes using the Penn World Tables as described below and in the appendix.

3. The empirical results

I now describe the results from estimating the Euler equations segmenting the countries according to the restrictions measures. I begin by describing some of the econometrics issues in the panel data estimation. I then turn to the results. Finally, I allow countries to face different real international interest rates. Interestingly, the basic results hold in all of these cases: the consumption growth rates in the restricted countries have a higher covariation with output growth than do the unrestricted countries.

3.1. Econometric issues

The data set includes observations for 72 countries over 27 years. Thus, the number of countries, J, equals 72 and the number of years, defined below as T, is 27. Since some countries have missing values for certain years, the total number of observations are 1762. [18] Due to the cross-sectional as well as time series dimension of the data set, the error terms are likely to be correlated both over time and across countries. To see how, consider the estimation equation (6), restated here for convenience:

$$\Delta c_{t+1}^j = \theta_0^j + \theta_0(t) + \beta^{\mathrm{R}} D(j,t)\,\Delta y_{t+1}^j + \beta^{\mathrm{U}}(1 - D(j,t))\,\Delta y_{t+1}^j + \epsilon_{t+1}^j.$$

Since aggregate consumption for individual countries tends to be serially correlated, it is likely that $\mathrm{E}(\epsilon_t^j \epsilon_{t-i}^j) \neq 0$ for $i = 1,2,\ldots,q$, where q is some maximum lag of autocorrelation. In addition, since countries may face similar shocks in a given year, it is likely that $\mathrm{E}(\epsilon_t^j \epsilon_t^{\ell}) \neq 0$ for $j \neq \ell$. Therefore, I allow for both possibilities in the estimation results.

In particular, I use a GMM estimator that allows for an MA(q) error process for each country and contemporaneous correlation in the error terms across countries. This estimator is the same as in Attanasio and Weber (1995), but it

[18] Due to these missing values, the data set is unbalanced as in Attanasio and Weber (1995). For expositional simplicity, I ignore this complication in the description of the estimator in the text. This issue is described in greater detail in the appendix.

allows for higher order moving average processes. [19] To see the form of the estimator, I define the stacked estimation equation (6) as: $\boldsymbol{C} = \boldsymbol{X}\delta + \epsilon$, where $\boldsymbol{C}$ is the $JT \times 1$ vector of stacked consumption growth rates, $\boldsymbol{X}$ is the $JT \times K$ matrix of stacked explanatory variables, δ is the $K \times 1$ parameter vector, and ϵ is the residual vector. Finally, let $\boldsymbol{Z}$ be the stacked $JT \times H$ instrumental variables matrix, where H is the number of instruments, and assume that the second moments of the error process converge to a positive definite matrix, Ω. Then, the estimator for δ is given by

$$\delta = \left[\boldsymbol{X}'\boldsymbol{Z}(\boldsymbol{Z}'\Omega\boldsymbol{Z})^{-1}\boldsymbol{Z}'\boldsymbol{X} \right]^{-1} \boldsymbol{X}'\boldsymbol{Z}(\boldsymbol{Z}'\Omega\boldsymbol{Z})^{-1}\boldsymbol{Z}'\boldsymbol{C}, \tag{7}$$

where the variance–covariance matrix of δ is

$$\mathrm{Var}(\delta) = \left[\boldsymbol{X}'\boldsymbol{Z}(\boldsymbol{Z}'\Omega\boldsymbol{Z})^{-1}\boldsymbol{Z}'\boldsymbol{X} \right]^{-1}. \tag{8}$$

In practice, estimators of these parameters and their variance–covariance matrix require identifying assumptions on the error process, ϵ, which, in turn impose zero restrictions on the Ω matrix. For this purpose, I make two identifying assumptions. First, I assume that $\mathrm{E}(\epsilon_t^j \epsilon_{t-q-\tau}^j) = 0$, for all $\tau > 0$. This assumption amounts to a lag length restriction on the error processes. I test for the appropriate lag length, q, using the Cumby and Huizinga (1992) 'L-test' described below.

The second identifying assumption I make follows Attanasio and Weber (1995). I assume that: $\mathrm{E}(\epsilon_t^j \epsilon_{t-\tau}^i) = 0$ for all $\tau > 0$, $i \neq j$. In other words, the error terms across countries are only correlated contemporaneously. Any time-series correlation across countries can be broken down into this contemporaneous correlation plus country-specific autocorrelation.

The appendix provides more details concerning the construction of this estimator.

3.2. Empirical results

Table 2 presents the results of estimating Eq. (6) using the GMM estimator described above. For instruments, I use the lagged right-hand side variables as well as lagged consumption growth following CM. I interact these instruments with the restrictions dummies so that the first stage regressions in the instrumental variables estimation are not constrained to be the same between restricted and unrestricted countries. Full lists of the instrumental variables are reported at the bottom of this and future tables.

Estimation requires a lag length assumption on the time series error processes, as described above. To get a sense for the appropriate lag length, I conduct a series of Cumby and Huizinga (1992) 'L-tests'. This test provides a χ^2 statistic of the

[19] Attanasio and Weber (1995) assume an MA(1) error process, while this estimator allows for an MA(q) where $q > 0$ by incorporating more autocovariance matrices for each country.

Table 2
L-tests and consumption Euler equations with common world interest rates [a]
$\Delta c^j_{t+1} = \theta^j_0 + \theta_0(t) + \beta^R D(j,t)\Delta y^j_{t+1} + \beta^U(1 - D(j,t))\Delta y^j_{t+1} + \epsilon^j_{t+1}$.

A. L-tests of hypothesis that ϵ^j_t is MA(q)

Sample period	$q =$	Proportion (number) of countries rejecting null at 5% marginal significance level for: 0	1	2
1967–1992		0.64	0.04	0.04
		(46)	(3)	(3)
1950–1992		0.69	0.07	0.04
		(50)	(5)	(3)

B. Consumption Euler equation estimates

Restriction	β^R	β^U	Marginal significance level H_0: $\beta^R = \beta^U$
Capital market	0.853 **	−0.094	< 0.001
	(0.054)	(0.080)	
Current accout	0.573 **	0.408 **	0.045
	(0.050)	(0.062)	
Export taxes	1.053 **	0.267 **	< 0.001
	(0.135)	(0.049)	
Import taxes	0.940 **	0.615 **	< 0.001
	(0.044)	(0.048)	
Import deposit requirements	1.076 **	0.004	< 0.001
	(0.045)	(0.072)	
Bilateral payments	0.504 **	0.457 **	0.577
	(0.068)	(0.050)	

[a] L-Tests are the tests from Cumby and Huizinga (1992) of the null hypothesis that the error term follows an MA(q) process. Panel A reports the proportion of the 72 countries for which the hypothesis is rejected at the 5% marginal significance level in favor of the hypothesis of a non-zero MA coefficient for one or more of the next 3 lags. The number of countries with rejections are given in parentheses. Panel B provides instrumental variables estimates and standard errors in parentheses allowing for contemporaneous correlation across countries and an MA(2) error process over time for each country. Instruments are lags of restricted output growth, unrestricted output growth, restricted consumption growth, and unrestricted consumption growth. The final column reports the marginal significance levels of a Wald test that $\beta^R = \beta^U$ and is distributed as $\chi^2(1)$.

** Significantly different from zero at the 95% confidence level.

hypothesis that $E(\epsilon_t \epsilon_{t-q-\tau}) = 0$, for $\tau > 0$ for a given country. Panel A reports the results for this test under the assumptions that $q = 0$ (white noise), 1, and 2. Since these tests are conducted for 72 countries for each value of q, I report the proportion of countries that reject the hypothesis at the 5% marginal significance level. If the processes are roughly similar, I would expect to reject the hypothesis when the hypothesis is true in approximately 5% of the countries. Thus, as a heuristic rule of thumb, I increase the lag length until roughly 5% or less of the countries reject the hypothesis.

Since the time-series dimension is short at only 27 years, this test may suffer from low power due to the small sample. To examine this possibility, the table also reports the tests for the full sample in the Summers–Heston data set from 1950 to 1992. However, the results do not show substantial differences in the implications for the order of the MA process. In this case, a lag length of $q = 2$ appears conservative.

Panel B reports the estimation results. Strikingly, for all six restrictions measures, $\beta^{\mathrm{R}} > \beta^{\mathrm{U}}$. The last column reports the marginal significance levels for the Wald test of the hypothesis that $\beta^{\mathrm{R}} = \beta^{\mathrm{U}}$. Since the hypothesis implies one linear restriction, the test statistic is distributed as $\chi^2(1)$. As the column shows, the hypothesis is rejected in all cases except for 'Bilateral Payments' at low marginal significance levels.

Thus, the evidence suggests that countries facing international restrictions are also likely to have a significantly higher income coefficient. According to the CM interpretation, this finding would suggest a higher proportion of liquidity constrained individuals.

3.3. Allowing for differences in international interest rates

The results above treat the international interest rate as the common expected consumption growth rate across countries by imposing fixed time effects. If countries are restricted from borrowing and lending in international markets, however, this interest rate may differ from the rate faced by unrestricted countries. In this section I examine the effects of allowing this rate to differ across countries. I assume that unrestricted countries have access to the U.S. dollar rate in real US terms. This assumption has commonly been used in the international literature on business cycles and the current account. [20]

When the US inflation rate is uncertain, the real interest rate is not known at time t. Assuming that R_t^{w} and consumption growth are joint log-normally distributed, the Euler equation can be rewritten

$$\Delta c_{t+1}^j = \left(\tfrac{1}{2}\gamma\sigma_j^2 + \tfrac{1}{2}(1/\gamma)\sigma_{Rj}^2 - \sigma_{RCj} + (1/\gamma)\ln(\rho^j)\right) + (1/\gamma)r_t^{\mathrm{w}} + \epsilon_{t+1}^j, \tag{9}$$

where now $\epsilon_{t+1}^j = (\Delta c_{t+1}^j - \mathrm{E}_t\Delta c_{t+1}^j) + (1/\gamma)(r_t^{\mathrm{w}} - \mathrm{E}_t r_t^{\mathrm{w}})$, the composite forecast error of consumption growth and the real interest rate. Also, σ_{Rj}^2 is the variance of the real interest rate and σ_{RCj} is the covariance between the interest rate and consumption growth.

Note that this equation has the same form as Eq. (6) except that the interpretation of the constant fixed effects changes and the interest rate is given by the

[20] For example, models in Baxter and Crucini (1995), Clarida (1990), and Obstfeld and Rogoff (1995) assume a world real interest rate.

Table 3
L-tests and consumption Euler equations with U.S. world interest rates [a]

$$\Delta c^j_{t+1} = \theta^j_0 + \theta^R_1 D(j,t) r^w_t + \theta^U_1 (1 - D(j,t)) r^w_t + \beta^R D(j,t) \Delta y^j_{t+1} + \beta^U (1 - D(j,t)) \Delta y^j_{t+1} + \epsilon^j_{t+1}$$

A. L-tests of hypothesis that ϵ^j_t is MA(q)

		Proportion (number) of countries rejecting null at 5% marginal significance level for:			
Sample period	$q =$	0	1	2	3
1967–1992		0.65	0.13	0.06	0.04
		(47)	(9)	(4)	(3)
1950–1992		0.61	0.08	0.01	0.05
		(44)	(6)	(1)	(4)

B. Consumption Euler equation estimates

Restriction	θ^R_1	θ^U_1	β^R	β^U	Marginal significance levels		
					H_0: $\theta^R_1 = \theta^U_1$, $\beta^R = \beta^U$	H_0: $\theta^R_1 = \theta^U_1$	H_0: $\beta^R = \beta^U$
Capital market	−0.035	−0.107 **	0.604 **	0.601 **	0.013	0.004	0.973
	(0.021)	(0.012)	(0.043)	(0.042)			
Current account	−0.044 **	−0.123 **	0.636 **	0.598 **	0.011	0.003	0.518
	(0.022)	(0.013)	(0.047)	(0.037)			
Export taxes	−0.016	−0.144 **	1.049 **	0.320 **	< 0.001	< 0.001	< 0.001
	(0.016)	(0.015)	(0.034)	(0.048)			
Import taxes	−0.029 *	−0.103 **	0.872 **	0.289 **	< 0.001	0.001	< 0.001
	(0.015)	(0.015)	(0.037)	(0.051)			
Import deposit requirements	−0.073 **	0.036 **	0.941 **	0.662 **	< 0.001	< 0.001	< 0.001
	(0.013)	(0.018)	(0.033)	(0.033)			
Bilateral payments	−0.058 **	−0.086 **	0.799 **	0.599 **	0.001	0.187	< 0.001
	(0.018)	(0.011)	(0.037)	(0.039)			

international US rate. In other words, this equation can be rewritten in the regression form:

$$\Delta c^j_{t+1} = \theta^j_0 + \theta_1 r^w_t + \beta X^j_t + \epsilon^j_{t+1}, \tag{10}$$

where now $\theta^j_0 = (\frac{1}{2}\gamma\sigma_j^2 + \frac{1}{2}(1/\gamma)\sigma^2_{Rj} - \sigma_{RCj} + (1/\gamma)\ln(\rho^j))$. I therefore estimate the Euler equation allowing the relationship with the logarithm of the common world interest rates, r^w_t, to vary according to whether countries are restricted or not:

$$\Delta c^j_{t+1} = \theta^j_0 + \theta^R_1 D(j,t) r^w_t + \theta^U_1 (1 - D(j,t)) r^w_t + \beta^R D(j,t) \Delta y^j_{t+1} + \beta^U (1 - D(j,t)) \Delta y^j_{t+1} + \epsilon^j_{t+1}. \tag{11}$$

If the international restrictions are important, the interest rate effects may differ: $\theta^R_1 \neq \theta^U_1$, for $i = 1,2$. When the Euler equation holds so that $\beta = 0$, the interest rate coefficient represents the elasticity of intertemporal substitution in consumption. However, empirical consumption studies have typically found this coefficient to be close to zero. [21] Also, under the CM interpretation, countries with a higher proportion of liquidity-constrained consumers would suggest the finding that $\beta^R > \beta^U$.

Table 3 reports the results of estimating Eq. (11) using the dollar London interbank offer rate (LIBOR) as the world nominal interest rate. This variable represents a known benchmark rate in international bond markets. The rate is adjusted for US inflation using the consumption price index from the Penn World Tables according to $r^w_t = \ln\{(1 + i^{us}_t) \times (P^{US}_t / P^{US}_{t+1})\}$ where i^{us}_t is the dollar LIBOR rate, and P^{US}_t is the US price index over the period. Since the ex post real interest rate is not known at time t, I use as instruments the lag of the right-hand

Notes to Table 3:

[a] L-Tests are the tests from Cumby and Huizinga (1992) of the null hypothesis that the error term follows an MA(q) process. Panel A reports the proportion of the 72 countries for which the hypothesis is rejected at the 5% marginal significance level in favor of the hypothesis of a non-zero MA coefficient for one or more of the next three lags. The number of countries with rejections are given in parentheses. Panel B provides instrumental variables estimates and standard errors in parentheses allowing for contemporaneous correlation across countries and an MA(3) error process over time for each country. Instruments are lags of restricted output growth, unrestricted output growth, restricted consumption growth, unrestricted consumption growth, the restricted U.S. real rate and the unrestricted U.S. real rate. The final three columns report the marginal significance levels of Wald tests that alternatively, $\theta^R_1 = \theta^U_1$ and $\beta^R = \beta^U$, $\theta^R_1 = \theta^U_1$, and $\beta^R = \beta^U$ and are distributed as $\chi^2(2)$, $\chi^2(1)$, and $\chi^2(1)$, respectively.

** Significantly different from zero at the 95% confidence level.

* Significantly different from zero at the 90% confidence level.

[21] See Hall (1988), for example.

side variables for restricted and unrestricted countries: the real interest rate, consumption growth rate, and income growth rate.

In Panel A, I report the results of the 'L-tests' for the lag length of the moving average error process. As before, the residuals show evidence of significant autocorrelation both over the sample period and an extended sample period dating back to 1950. A conservative assumption therefore appears to be that the error process is MA(3).

Panel B gives the estimation results. As in domestic studies, the interest rate coefficients, θ_1^i, are close to zero, but are also negative, apparently due to a negative correlation between the US price index and the consumption growth rate. [22] In all cases but 'Bilateral Payments' these interest coefficients for restricted and unrestricted countries are significantly different from each other.

Interestingly, the income coefficients show the same pattern as before. In all cases, either $\beta^{\mathrm{R}} > \beta^{\mathrm{U}}$ or else the coefficients are insignificantly different as for the broader measures such as 'Capital Market' and 'Current Account'. Generally, these results corroborate those in Table 2, finding that the coefficients for restricted countries exhibit greater income sensitivity.

3.4. Allowing for country-specific real exchange rate changes

An important underlying assumption in the estimates above is that the real international interest rate is common to all countries. However, even if international capital markets are open to all countries and the countries face the same nominal interest rate, real interest rates will differ if real exchange rates are not constant over time. Thus, if i_t is the one-year nominal interest rate in dollars observed at time t, the ex post real interest rate over period t to $t+1$ for country j is

$$R_t^j = (1+i_t)\left(S_{t+1}^j P_t^j / S_t^j P_{t+1}^j\right), \tag{12}$$

where S_t^j is the price of dollars in terms of currency j and P_t^j is the price level of country j both at time t. Under purchasing power parity, the real exchange rate, $P_t^j/(S_t^j P_t^{\mathrm{us}})$, is constant so that the real interest rate is common to all countries as assumed in Table 3. A substantial amount of research has demonstrated that purchasing power parity is rejected, however. [23] Therefore, even a common nominal world interest rate may imply substantially different domestic real interest rates.

I therefore test for differences between liquidity constraints among the re-

[22] Omitting the nominal LIBOR rate and running the same estimates on the US inflation rate alone gives a similar negative coefficient. This negative covariation does not show up in domestic studies since cross-country relationships are not typically analyzed.

[23] For a recent survey, see Froot and Rogoff (1995).

stricted and unrestricted countries by conditioning on country-specific interest rates in the following way:

$$\Delta c_{t+1}^{j} = \theta_0^{j} + \theta_1^{R} D(j,t) r_{t+1}^{j} + \theta_1^{U}(1 - D(j,t)) r_{t+1}^{j} + \beta^{R} D(j,t) \Delta y_t^{j} + \beta^{U}(1 - D(j,t)) \Delta y_t^{j} + \epsilon_{t+1}^{j}, \qquad (13)$$

and testing the hypotheses: $\theta_1^R = \theta_1^U$, $\beta^R = \beta^U$. For the interest rates, I use the LIBOR rate adjusted by the country-specific consumption price indices in the Penn World Tables. The appendix describes the construction of these series. Since the real interest rate is not known at time t, I use as instruments the lagged real interest rates, lagged income growth rates, and lagged consumption growth rates for restricted and unrestricted countries.

Table 4 reports the results of estimating these equations. Panel A summarizes the 'L-test' results, suggesting an MA(3) process as a conservative assumption. Panel B shows the estimates of Eq. (13). These estimates are strikingly similar to the results using common world interest rates in Table 3, Panel B. All measures but the two broadest restriction measures reject the hypothesis that the parameters are common across restricted and unrestricted countries. The interest rate coefficients are small but negative and, in half the cases, the hypothesis that: $\theta_1^U = \theta_1^R$ is rejected. Finally, when the income coefficients are significantly different, $\beta^R > \beta^U$. Overall, the results appear to corroborate and strengthen the evidence in Table 3.

4. Alternative explanations

The preceding discussion has treated the positive covariation between expected consumption growth and expected output growth as evidence against intertemporal consumption smoothing. In this section, I examine four alternative explanations for this finding even when the intertemporal consumption Euler equation holds. These explanations are based upon various types of omitted variables that are correlated with income growth. While these explanations are by no means exhaustive, they are among the most often-cited alternative interpretations for Euler equation rejections.

The four explanations I examine are: (1) measurement error in income growth and/or consumption growth, (2) conditional heteroscedasticity in consumption growth, (3) habit persistence, and (4) non-separabilities in utility between traded and non-traded goods or leisure. In the case of heteroscedasticity, I present evidence suggesting that the basic conclusions above are maintained when allowing for time-varying variances.

4.1. Measurement error

To show how measurement error can explain the above results, I make several simplifying assumptions. First, I assume that the interest rate is constant and is therefore captured by the constant term. In this case, the estimation equation can

Table 4

L-tests and consumption Euler equations with time-varying world interest rates differing by real exchange rates [a]

$$\Delta c^j_{t+1} = \theta^j_0 + \theta^R_1 D(j,t) r^j_t + \theta^U_1 (1 - D(j,t)) r^j_t + \beta^R D(j,t) \Delta y^j_{t+1} + \beta^U (1 - D(j,t)) \Delta y^j_{t+1} + \epsilon^j_{t+1}$$

A. *L*-tests of hypothesis that ϵ^j_t is MA(q)

		Proportion (number) of countries rejecting null at 5% marginal significance level for:			
Sample period	$q =$	0	1	2	3
1967–1992		0.63	0.10	0.03	0.04
		(45)	(7)	(2)	(3)
1950–1992		0.64	0.13	0.08	0.04
		(46)	(9)	(6)	(3)

B. Consumption Euler equation estimates

Restriction	θ^R_1	θ^U_1	β^R	β^U	Marginal significance levels		
					H_0: $\theta^R_1 = \theta^U_1$, $\beta^R = \beta^U$	H_0: $\theta^R_1 = \theta^U_1$	H_0: $\beta^R = \beta^U$
Capital market	−0.024	−0.069 **	0.606 **	0.628 **	0.140	0.56	0.722
	(0.020)	(0.012)	(0.020)	(0.012)			
Current account	−0.052 **	−0.074 **	0.633 **	0.624 **	0.690	0.391	0.891
	(0.022)	(0.012)	(0.047)	(0.036)			
Export taxes	−0.023	−0.095 **	1.044 **	0.354 **	< 0.001	0.002	< 0.001
	(0.016)	(0.015)	(0.034)	(0.048)			
Import taxes	−0.024 *	−0.026 **	0.874 **	0.318 **	< 0.001	0.914	< 0.001
	(0.016)	(0.013)	(0.037)	(0.049)			
Import deposit requirements	−0.062 **	0.054 **	0.941 **	0.667 **	< 0.001	< 0.001	< 0.001
	(0.015)	(0.016)	(0.033)	(0.032)			
Bilateral payments	−0.055 **	−0.046 **	0.798 **	0.622 **	0.002	0.686	< 0.001
	(0.019)	(0.011)	(0.037)	(0.039)			

be written as in Eq. (6) where the time effect is now constant. Second, I treat the expected growth rate as observable with error at time t, defined as $\phi_t^j = \mathrm{E}_t \Delta y_{t+1}^j + u_t^j$, where u is measurement error. In this case, the estimation equation can be written as

$$\Delta c_{t+1}^j = \theta_0^j + \beta^{\mathrm{R}} D(j,t)\phi_t^j + \beta^{\mathrm{U}}(1 - D(j,t))\phi_t^j + \xi_{t+1}^j, \tag{6'}$$

where $\xi_{t+1}^j \equiv \epsilon_{t+1}^j + \beta^{\mathrm{R}} D(j,t) u_t^j + \beta^{\mathrm{U}}(1 - D(j,t)) u_t^j$. Third, I consider the effects of running an OLS regression of consumption growth on the observed proxy for expected output growth. Clearly, the actual estimation using instrumental variables is more complicated, but retains the basic features.[24] Under these assumptions, the probability limit of β^{i}, for $i = \mathrm{R,U}$, under perfect consumption smoothing is[25]

$$\operatorname{plim}(\beta^i) = \operatorname{Cov}(\Delta c_{t+1}^j, \phi_t^j)/\operatorname{Var}(\phi_t^j) = \operatorname{Cov}(\Delta c_{t+1}^j, u_t^j)/\operatorname{Var}(\phi_t^j), \tag{14}$$

where the second equality follows since: $\operatorname{Cov}(\Delta c_{t+1}^j, \mathrm{E}_t \Delta y_{t+1}^j) = 0$. Moreover, with time-aggregation problems, it is likely that the measurement error in expected consumption will be correlated with consumption growth. Thus, even when the Euler equation holds, time-aggregation may lead to the finding that $\beta^i > 0$. Although this discussion focuses upon measurement error in income growth, this correlation is likely to be exacerbated by measurement error in consumption growth as well.

Notes to Table 4:

[a] L-tests are the tests from Cumby and Huizinga (1992) of the null hypothesis that the error term follows an MA(q) process. Panel A reports the proportion of the 72 countries for which the hypothesis is rejected at the 5% marginal significance level in favor of the hypothesis of a non-zero MA coefficient for one or more of the next three lags. The number of countries with rejections are given in parentheses. Panel B provides instrumental variables estimates and standard errors in parentheses allowing for contemporaneous correlation across countries and an MA(3) error process over time for each country. Instruments are lags of restricted output growth, unrestricted output growth, restricted consumption growth, unrestricted consumption growth, the restricted U.S. real rate and the unrestricted U.S. real rate. The final three columns report the marginal significance levels of Wald tests that alternatively, $\theta_1^{\mathrm{R}} = \theta_1^{\mathrm{U}}$ and $\beta^{\mathrm{R}} = \beta^{\mathrm{U}}$, $\theta_1^{\mathrm{R}} = \theta_1^{\mathrm{U}}$, and $\beta^{\mathrm{R}} = \beta^{\mathrm{U}}$ and are distributed as $\chi^2(2)$, $\chi^2(1)$, and $\chi^2(1)$, respectively.

** Significantly different from zero at the 95% confidence level.

* Significantly different from zero at the 90% confidence level.

[24] Weak instruments can also introduce biases in β as has been noted by Campbell and Mankiw (1990) as well as others. In the present case, lagged variables appear to have significant explanatory power for the right-hand side variables.

[25] This result follows for OLS since the information matrix is block diagonal between β^{R} and β^{U}.

Note that to explain the finding that $\beta^{R} > \beta^{U}$, this story suggests that:

$$\begin{aligned}\text{plim}(\beta^{R}) &= \text{Cov}(\Delta c_{t+1}^{R}, u_t^{R})/\text{Var}(\phi_t^{R}) > \text{plim}(\beta^{U}) \\ &= \text{Cov}(\Delta c_{t+1}^{U}, u_t^{U})/\text{Var}(\phi_t^{U}).\end{aligned}$$

Thus, the finding would require a higher covariation between measurement error for the restricted countries than the unrestricted countries along with a lower variance of the expected income growth proxy.

There is no immediately apparent reason for this pattern to hold. Note, however, that this pattern could hold if measurement error between consumption and income were correlated and this measurement error were serially correlated over time to a greater degree for restricted countries than for unrestricted countries. Without more information about differences in patterns of measurement error, this explanation is difficult to assess.

4.2. Conditional heteroscedasticity

Another potential explanation for finding $\beta \neq 0$ in the above regressions is conditional heteroscedasticity in consumption growth. To see how, assume that consumption growth is conditionally log normal. Replacing the variance in Eq. (2) with its conditional version gives

$$\Delta c_{t+1}^{j} = \left(\tfrac{1}{2}\text{Var}_t(\Delta c_{t+1}^{j}) + (1/\gamma)\ln(\rho_j)\right) + (1/\gamma) r_t^{w} + \epsilon_{t+1}^{j}. \tag{15}$$

As noted by Zeldes (1989), covariation between σ_{jt}^2 and the right-hand side variables such as income growth can bias the regression coefficient, similar to the measurement error problem noted above. [26]

One way to examine whether the findings that $\beta^{i} > 0$ and $\beta^{R} > \beta^{U}$ are due to underlying covariation between consumption growth and its variance would be to identify and estimate a variance time series process such as GARCH. However, the relatively short time dimension of the data set precludes estimating these series for each country. Instead, I evaluate the possible effects of time variation in the variance by conditioning consumption growth on a noisy measure of this variance. In particular, I assume that the conditional variance of the consumption process is proportional to the squared ex post residuals plus a forecast error. [27] In other words,

$$\text{Var}_t(\Delta c_{t+1}^{j}) = \alpha(\epsilon_{t+1}^{j})^2 + v_{t+1}, \tag{16}$$

where $\alpha > 0$. I focus upon the more parsimonious case where the interest rate is the common fixed time effect since the results were little affected by using the

[26] More recently, other authors such as Carroll (1996) have noted that the effects of variation in consumption variance can affect precautionary savings, altering the interpretation of β.

[27] See Lewis (1991) for a more detailed description of this approach.

Table 5
Consumption Euler equation with time-varying variances [a]
$\Delta c^j_{t+1} = \theta^j_0 + \theta_0(t) + \alpha\sigma^j_t + \beta^R D(j,t)\Delta y^j_{t+1} + \beta^U(1-D(j,t))\Delta y^j_{t+1} + \epsilon^j_{t+1}$

Restriction	α	β^R	β^U	Marginal significance level H_0: $\beta^R = \beta^U$
Capital market	9.009	1.036 **	−0.218	< 0.001
	(1.036)	(0.123)	(0.174)	
Current account	6.406	0.886 **	0.009	< 0.001
	(1.169)	(0.103)	(0.178)	
Export taxes	5.253	1.220 **	0.059	0.016
	(1.245)	(0.454)	(0.142)	
Import taxes	8.890	1.405 **	0.924 **	< 0.001
	(0.448)	(0.067)	(0.059)	
Import deposit	8.723	0.817 **	−0.074	< 0.001
requirements	(1.275)	(0.102)	(0.187)	
Bilateral	3.932	0.894 **	0.330 **	< 0.001
payments	(0.517)	(0.093)	(0.112)	

[a] L-tests are the tests from Cumby and Huizinga (1992) of the null hypothesis that the error term follows an MA(q) process. The table provides instrumental variables estimates and standard errors in parentheses allowing for contemporaneous correlation across countries and an MA(2) error process over time for each country. Instruments are lags of restricted output growth, unrestricted output growth, restricted consumption growth, and unrestricted consumption growth. The final column reports the marginal significance levels of a Wald test that $\beta^R = \beta^U$ and is distributed as $\chi^2(1)$. Standard errors under α do not correct for the generated regressors problem and therefore likely understate the true variance of the estimate.

** Significantly different from zero at the 95% confidence level.

actual interest rates. Thus, substituting the observed squared residual for the conditional variance in Eq. (6), the regression equation can be rewritten as

$$\Delta c^j_{t+1} = \theta^j_0 + \theta_0(t) + \alpha\left(\epsilon^j_{t+1}\right)^2 + \beta^R D(j,t)\Delta y^j_{t+1} + \beta^U(1-D(j,t))\Delta y^j_{t+1} + \xi^j_{t+1} \tag{17}$$

where now $\theta^j_0 = (1/\gamma)\ln(\rho^j)$, $\xi^j_{t+1} \equiv \epsilon^j_{t+1} + v^j_{t+1}$, and the squared residual provides a noisy measure of the variance. [28] Since the ex post variance is not known at time t, the estimation must use instruments in the time t information set. As before, I use the lagged right-hand side variables.

Table 5 reports the results of estimating these equations for the six restrictions measures and assuming an MA(2) error process as in Table 2. I adopt a two step procedure for estimation using the measure of conditional variance. First, I estimate the consumption equations with instrumental variables as in Table 2,

[28] Dynan (1993) takes a similar approach with household data. She averages over squared quarterly consumption within a year, finding that the precautionary savings effect is small. The basic conclusions in the text are preserved when using squared total consumption instead of squared unexpected consumption.

saving the residuals. Second, I treat the squared residuals as regressors in Eq. (17) to provide measures of the conditional variance and re-estimate the equations with instrumental variables. I make no correction for the 'generated-regressors' problem created by the first-stage estimation so that the standard errors must be viewed with caution. Nevertheless, conditioning consumption growth on measures of the variance potentially absorbs the variation due to σ_{jt}. Therefore, if the finding that $\beta^{R} > \beta^{U} > 0$ disappears when controlling for the variance, it would seem likely this bias is largely responsible for the result.

Table 5 suggests controlling for this variance does not overturn the basic conclusions, however. The basic pattern of higher coefficients for the restricted countries is preserved. Since the standard errors do not correct for the first-stage regression, technically speaking the hypothesis tests cannot be conducted. However, to get a feel for whether these hypotheses would be rejected based solely upon the second stage regressions, I report these statistics. Interestingly, while the hypothesis that $\beta^{R} = 0$ is rejected, the same hypothesis for $\beta^{U} = 0$ can be rejected only for two of the finer measures, 'Import Taxes' and 'Bilateral Payments'. Overall, the basic results do not appear to be driven by omitting variation in the consumption variance.

4.3. Habit persistence

Utility functions with habit persistence can also introduce omitted variables that may be correlated with variables in the lagged information set. Constantinides (1990) and others have argued that the utility from current consumption depends upon how this consumption compares to past consumption levels. To relate this utility function to the estimation above, it is convenient to use of the form of this utility function specified by Abel (1990): [29]

$$u(C_t) = (C_t / C_{t-1}^{\eta})^{1-\gamma} / (1-\gamma). \tag{18}$$

As Abel explains, η represents the habit persistence parameter where $\eta = 0$ implies the standard time-additive iso-elastic utility case.

Substituting this utility function into the Euler equation (1), and taking the logarithm implies:

$$\Delta c_{t+1}^{j} = \left(\tfrac{1}{2}\sigma_j^2 + (1/\gamma)\ln(\rho^j)\right) + (1/\gamma) r_t^{\mathrm{w}} + (\eta/\gamma)\Delta c_t^j + \epsilon_{t+1}^j. \tag{19}$$

Relating this Euler equation to the estimation equation (6) shows that, if expected income growth is positively correlated with current consumption growth, then we will find $\beta > 0$ even when the true $\beta = 0$. Thus, habit persistence can explain a non-zero coefficient on expected income growth.

[29] Abel (1990) introduces a more general utility form that simultaneously allows for habit persistence and 'keeping up with the Joneses', a preference for individual consumption levels relative to the aggregate consumption level. In the text, I focus only upon the former effect.

On the other hand, the finding that $\beta^{R} > \beta^{U}$ requires an additional feature beyond habit persistence. For restricted relative to unrestricted countries, either the covariance between expected income growth and lagged consumption must be higher or else the degree of habit persistence, η, must be higher.

4.4. *Non-separabilities in utility between tradeables and non-tradeables*

Finally, I analyze the potential biases introduced by non-separabilities in utility between tradeables and non-tradeables. In primarily cross-sectional analysis, Lewis (1996) shows that non-separabilities are important for understanding the lack of risk-sharing internationally in the presence of capital market restrictions. Unfortunately, the data set does not include enough time-series observations to analyze the significance of non-tradeables in intertemporal consumption smoothing.

Nevertheless, if tradeables and non-tradeables are complements as recent studies have assumed, then a higher expenditure share on non-tradeables by restricted countries would also be consistent with the findings above. [30] This share may differ across countries due to factors such as government spending and services. Also, the basic reasoning applies for leisure when labor is immobile internationally. [31] I sketch out this result below and relegate the details to the appendix. I first treat the aggregate consumption measure as tradeable consumption to illustrate the basic point, and later discuss the implications when aggregate consumption is a composite of both tradeables and non-tradeables.

Assume that instantaneous utility is a function of both tradeable consumption and non-tradeables, defined as N_t^j for country j. Specifically, suppose the form of this utility function is

$$U(C_t^j, N_t^j) = \Psi(C_t^j, N_t^j)^{(1-\gamma)}/(1-\gamma), \tag{20}$$

where Ψ is a linearly homogeneous function. A standard assumption in the international business cycle literature is that borrowing in international markets requires payment in terms of the tradeables good. Thus, the Euler equation becomes

$$\mathrm{E}_t\left[\,\rho^j U_C(C_{t+1}^j, N_{t+1}^j)/U_C(C_t^j, N_t^j)\right] = 1/R_t^{\mathrm{w}}, \tag{21}$$

where U_C is the marginal utility with respect to the tradeable consumption. Furthermore,

$$U_C(C_t^j, N_t^j) = \Psi(C_t^j, N_t^j)^{-\gamma}\Psi_C(C_t^j, N_t^j).$$

Then, assuming that tradeables and non-tradeables consumption growth are joint

[30] See the summary of the literature described in Baxter et al. (1995), for instance.

[31] Baxter and Jermann (1994) show that non-separabilities between consumption and leisure can explain the Campbell and Mankiw (1989, 1991) result even when individuals are not constrained.

log-normally distributed, the appendix shows that the logarithm of the Euler Eq. (21) can be written as

$$\Delta c_{t+1}^{j} = \theta_0^j + \theta_1 r_t^{w} + \alpha \mathrm{E}_t \Delta n_{t+1}^{j} + \epsilon_{t+1}^{j}, \tag{22}$$

where now $\theta_0^j \equiv (\sigma_{\psi j} + \ln(\rho^j))\theta_1$, $\theta_1 \equiv (\gamma(1 - x_{\mathrm{N}}) + (x_{\mathrm{N}} \zeta))^{-1}$, $\sigma_{\psi j}$ is the variance of the growth rate in the marginal utility with respect to tradeables, x_{N} is the expenditure share on non-tradeables, and ζ is the inverse of the elasticity of substitution between tradeables and non-tradeables.

More importantly, tradeables consumption growth depends upon expected non-tradeables consumption growth according to the parameter α where: $\alpha \equiv (\zeta - \gamma)/(\zeta + \gamma(1 - x_{\mathrm{N}})x_{\mathrm{N}}^{-1})$. Thus, $\alpha > 0$ as long as $\zeta > \gamma$; i.e., as long as tradeables and non-tradeables are net complements. The intuition behind this result is straightforward. If tradeables and non-tradeables are net complements, then lower expected future non-tradeables consumption will increase the marginal utility of future non-tradeables and, hence, tradeables consumption. Therefore, consumers will try to move current tradeables consumption into future periods, inducing a positive relationship between tradeables and non-tradeables consumption growth.

Consider now the implications for the estimation of Eq. (6) above rewritten here:

$$\Delta c_{t+1}^{j} = \theta_0^j + \theta_0(t) + \beta^{\mathrm{R}} D(j,t) \Delta y_{t+1}^{j} + \beta^{\mathrm{U}}(1 - D(j,t)) \Delta y_{t+1}^{j} + \epsilon_{t+1}^{j}.$$

Since non-tradeables consumption and non-tradeables output are equal by definition, then the covariance between expected output growth and expected non-tradeables growth is likely positive: $\mathrm{Cov}(\mathrm{E}_t \Delta y_{t+1}^{j}, \mathrm{E}_t \Delta n_{t+1}^{j}) > 0$. As a result, the estimation is likely to find $\beta^i > 0$ even in the presence of perfect intertemporal smoothing of tradeables consumption.

Consider next the pattern found above that $\beta^{\mathrm{R}} > \beta^{\mathrm{U}}$. Conventional wisdom suggests that developing countries tend to have a higher share of consumption spending on non-tradeables. If developing countries are more likely to face international restrictions, as the evidence in Table 1 suggests, then complementarity between tradeables and non-tradeables would imply this pattern.

To see why, suppose that the covariance between non-tradeables and output are the same between restricted and non-restricted countries. Clearly, even in this case, the coefficients, β^i, will differ across restricted and unrestricted countries if the expenditure share on non-tradeables is higher for restricted countries. Specifically, recall that the coefficient on expected non-tradeables growth is $\alpha \equiv (\zeta - \gamma)/(\zeta + \gamma(1 - x_{\mathrm{N}})x_{\mathrm{N}}^{-1})$, where x_{N} is the expenditure share on non-tradeables. By inspection, it is clear that countries with higher non-tradeables expenditures have higher coefficients. Thus, if $\alpha^{\mathrm{R}} > \alpha^{\mathrm{U}}$, we would expect to find $\beta^{\mathrm{R}} > \beta^{\mathrm{U}}$, as above.

Of course, the evidence in this paper is based upon aggregate consumption as the regressor, a variable that combines tradeables and non-tradeables. Nevertheless, this analysis suggests the direction of the effect on β from complementarity between tradeables and non-tradeables.

5. Concluding remarks

This paper has examined the differences in consumption sensitivity to output growth across various measures of international restrictions. Campbell and Mankiw (1989, 1991) have suggested that greater sensitivity of expected consumption to expected output growth may be interpreted as a higher proportion of individuals who are liquidity constrained. While some of the restrictions examined do not necessarily imply capital market restrictions per se, governments who impose restrictions are more likely to impose other types of regulations. Thus, the approach of this paper has been to let the data speak on whether consumers in countries facing six different international restrictions have significantly different income sensitivities.

Interestingly, the evidence in this paper suggests that these restrictions matter. Specifically, across a broad range of restrictions measures and specifications for the interest rate, the income sensitivity is significantly higher for restricted countries than unrestricted countries. These differences are most pronounced for the restrictions that affect a smaller proportion of the world's countries. Moreover, the relationship remains when allowing for precautionary savings due to time-variation in consumption variances.

The paper also provides an analysis of alternative explanations for the results including differences across restricted and unrestricted countries in: (a) the patterns of measurement error, (b) habit persistence parameters, and/or (c) expenditure shares on tradeables and non-tradeables goods when these goods are complements in utility. Future research should examine whether these alternative explanations are important.

Acknowledgements

For useful comments and suggestions, I am grateful to Chris Carroll, Francesco Giavazzi, Nick Souleles, Guglielmo Weber, two anonymous referees, and participants at the International Seminar on Macroeconomics, Columbia University, the New York Federal Reserve Bank, the Board of Governors of the Federal Reserve, and the Wharton Macro Lunch Group. Any errors are mine alone.

Appendix A. Data sources and construction

A.1. Consumption and output data

Following standard practice in the literature, the consumption and income data were taken from the Penn World Tables described in Summers and Heston (1991), updated using the most recent data available in the Mark 5.6 version. Specifically, output is the series RGDPCH 'Real GDP per capital in constant dollars' using 1985 as the base year. Consumption is series C, the share of GDP times output. Countries with missing values for a given year are dropped from estimation, including calculation of the fixed time effects. In choosing the countries, I follow Obstfeld (1994a) and Tesar (1995) in taking all countries with data quality of C- or better. These are: Argentina, Australia, Austria, Barbados, Belgium, Bengaladesh, Bolivia, Botswana, Brazil, Cameroon, Canada, Chile, Columbia, Costa Rica, Cyprus, Denmark, Dominican Republic, Ecuador, El Salvador, Finland, France, Germany, Greece, Guatemala, Honduras, Hong Kong, Hungary, Iceland, India, Indonesia, Iran, Ireland, Israel, Italy, Ivory Coast, Jamaica, Japan, Kenya, Luxembourg, Malaysia, Mexico, Morroco, Netherlands, New Zealand, Norway, Pakistan, Panama, Paraguay, Peru, Philippines, Poland, Portugal, Senegal, Singapore, South Africa, South Korea, Spain, Sri Lanka, Sweden, Switzerland, Syria, Tanzania, Thailand, Trinidad and Tobago, Tunisia, Turkey, United Kingdom, United States, Uruguay, Venezuela, Yugoslavia, and Zimbabwe.

A.2. Interest Data

The world nominal interest rate data are the LIBOR rate from the Bank of England's *Quarterly Bulletin*. To adjust for inflation, this series is multiplied by the ratio of the $t+1$ domestic price at international prices to the same price at time t. This price is calculated by taking the ratio of future consumption at international prices to the 1985 real consumption level.

A.3. Capital Market Restrictions Data

The data on capital market restrictions are from the summary tables at the end of *The Annual Report on Exchange Arrangements and Exchange Restrictions* (1967 through 1993). These data equal one if there was a restriction during the year, and a zero otherwise. See the appendix in Lewis (1996) for a more detailed discussion of some of the restrictions measures.

Appendix B. Estimation

As described in the text, the GMM estimator is given by

$$\delta = \left[X'Z(Z'\Omega Z)^{-1}Z'X\right]^{-1} X'Z(Z'\Omega Z)^{-1}Z'C$$

where the variance–covariance matrix of δ is

$$\mathrm{Var}(\delta) = \left[X'Z(Z'\Omega Z)^{-1}Z'X\right]^{-1}.$$

This estimation requires an estimate of Ω. I first estimate the stacked equation: $C = X\delta + \epsilon$ using instrumental variables, Z, to obtain a consistent estimate of ϵ. I then follow Attanasio and Weber (1995) by using these residuals to construct an estimate of $Z'\Omega Z$ according to

$$Z'\Omega Z = A_0 + A_1 + aA_2$$

where

$$A_0 = \sum_{j=1}^{J} B_{jj} \sum_{t=1}^{T_{jj}} z_{t*}^{j\prime} z_{t*}^{j} \left(\epsilon_t^j\right)^2,$$

$$A_1 = \sum_{\tau=1}^{q} \left[(1-\tau)/(1+q)\right] \sum_{j=1}^{J} B_{jj} \sum_{t=\tau+1}^{T_{jj}} z_{t*}^{j\prime} z_{t-\tau*}^{j} \epsilon_t^j \epsilon_{t-\tau}^j,$$

$$A_2 = \sum_{\ell=1, \ell=j}^{J} \sum_{j=1}^{J} B_{\ell j} \sum_{t=1}^{T_{\ell j}} z_{t*}^{j\prime} z_{t*}^{\ell} \epsilon_t^j \epsilon_t^{\ell},$$

where $T_{j\ell}$ is the number of observations in common for countries j and ℓ, ϵ_t^j is defined as the residual for country j at time t, z_{t*}^j is the $H \times 1$ row vector of instruments of country j at time t, and $B_{j\ell}$ is the inverse of the total number of observations in each subgroup of j and ℓ. Note that A_1 is the sum over the covariance matrices of the countries using a Newey and West (1987) estimator for autocovariances up to lag q, assuming an error process of MA(q). A_2 captures the cross-sectional covariation at time t across countries. As described in Attanasio and Weber (1995), a is an ad hoc weight less than one. The Wald tests in the tables are constructed using the variance–covariance matrix given by Eq. (8).

Appendix C. Derivation of relationship between tradeables and non-tradeables growth

Recall that the utility function is given by $U(C_t^j, N_t^j) = \Psi(C_t^j, N_t^j)^{(1-\gamma)}/(1-\gamma)$. Then substituting the marginal utility with respect to tradeables, $U_C(C_t^j, N_t^j) = \Psi(C_t^j, N_t^j)^{-\gamma}\Psi_C(C_t^j, N_t^j)$, into the Euler equation (21) gives

$$\begin{aligned} &\mathrm{E}_t\left[\rho^j \Psi\left(C_{t+1}^j, N_{t+1}^j\right)^{-\gamma} \Psi_C\left(C_{t+1}^j, N_{t+1}^j\right) / \Psi\left(C_t^j, N_t^j\right)^{-\gamma} \Psi_C\left(C_t^j, N_t^j\right)\right)\right] \\ &\quad = 1/R_t^{\mathrm{w}}. \end{aligned} \tag{C.1}$$

Assuming joint log-normality and taking the natural logarithm of Eq. (C.1) implies:

$$\begin{aligned} -\gamma \mathrm{E}_t \Delta \ln\Psi\left(C_{t+1}^j, N_{t+1}^j\right) + \Delta \mathrm{E}_t \ln\Psi_C\left(C_{t+1}^j, N_{t+1}^j\right) = -\sigma_{\psi j}^2 - \ln\left(\rho^j\right) \\ - r_t^{\mathrm{w}}. \end{aligned} \tag{C.2}$$

where $\sigma_{\psi j}^2 \equiv \frac{1}{2}\gamma^2 \operatorname{Var}\Delta\Psi + \frac{1}{2}\operatorname{Var}\Delta\Psi_C - \gamma \operatorname{Cov}(\Delta\Psi, \Delta\Psi_C)$. By the linear homogeneity of Ψ, the following linear approximations hold:

$$\Delta \ln\Psi\left(C_{t+1}^j, N_{t+1}^j\right) = \left(1 - x_{\mathrm{N}}\right) \Delta c_{t+1}^j + x_{\mathrm{N}} \Delta n_{t+1}^j \tag{C.3}$$

and

$$\Delta \ln\Psi_C\left(C_{t+1}^j, N_{t+1}^j\right) = -x_{\mathrm{N}} \zeta\left(\Delta c_{t+1}^j - \Delta n_{t+1}^j\right), \tag{C.4}$$

where ζ is the inverse of the elasticity of substitution between tradeables and nontradeables consumption growth. Substituting Eqs. (C.3) and (C.4) into Eq. (C.2) and actual for expected Δc_{t+1}, yields Eq. (22) in the text.

References

Abel, A.B., 1990, Asset prices under habit formation and catching up with the Joneses, American Economic Review 80, 38–42.

Attanasio, O.P. and G. Weber, 1995, Is consumption growth consistent with intertemporal optimization? Evidence from the Consumer Expenditure Survey, Journal of Political Economy 103, 1121–1157.

Backus, D.K., P.J. Kehoe and F.E. Kydland, 1992, International real business cycles, Journal of Political Economy 100, 745–775.

Baxter, M. and M. Crucini, 1995, Business cycles and the asset structure of foreign trade, International Economic Review, 821–854.

Baxter, M. and U.J. Jermann, 1994, Household production and the excess sensitivity of consumption to current income, Working paper (University of Virginia, Charlottesville, VA).

Baxter, M. and U.J. Jermann, 1995, The international diversification puzzle is worse than you think, Working paper no. 5019 (National Bureau of Economic Research, Cambridge, MA).

Baxter, M., U.J. Jermann and R.G. King, 1995, Nontraded goods, nontraded factors, and international non-diversification, Working paper no. 5175 (National Bureau of Economic Research, Cambridge, MA).

Bayoumi, T., 1993, Financial deregulation and consumption in the United Kingdom, The Review of Economics and Statistics 75, 536–539.

Campbell, J.Y. and G. Mankiw, 1989, Consumption, income, and interest rates: Reinterpreting the time series evidence, In: O.J. Blanchard and S. Fischer, eds., NBER macroeconomics annual, 1989 (MIT Press, Cambridge, MA).

Campbell, J.Y. and G. Mankiw, 1990, Permanent income, current income, and consumption, Journal of Business and Economic Statistics 8, 265–297.

Campbell, J.Y. and G. Mankiw, 1991, The response of consumption to income: A cross-country investigation, European Economic Review 35, 723–756.

Carroll, C.D., 1996, Buffer stock saving and the life cycle/permanent income hypothesis, Quarterly Journal of Economics, forthcoming.

Clarida, R.H., 1990, International lending and borrowing in a stochastic stationary equilibrium, International Economic Review 31, 180–202.

Constantinides, G., 1990, Habit formation: A resolution of the equity premium puzzle, Journal of Political Economy 98, 519–543.

Cumby, R.E. and J. Huizinga, 1992, Testing the autocorrelation structure of disturbances in ordinary least squares and instrumental variables regressions, Econometrica 60, 185–195.

Deaton, A., 1992, Understanding consumption (Oxford University Press, New York).

Devereux, M.B., A.W. Gregory and G.W. Smith, 1992, Realistic cross-country consumption correlations in a two-country, equilibrium business cycle model, Journal of International Money and Finance 11, 3–16.

Dynan, K., 1993, How prudent are consumers? Journal of Political Economy 101, 1104–1113.

Folkerts-Landau, D. and T. Ito, eds., 1995, International capital markets: Developments, prospects, and policy issues (Ch. I and Ch. VII in Vol. I: Background papers) (International Monetary Fund, Washington, DC).

French, K.R. and J.M. Poterba, 1991, Investor diversification and international equity markets, American Economic Review 81, 222–226.

Froot, K. and K. Rogoff, 1995, Perspectives on PPP and long-run real exchange rates, In: G. Grossman and K. Rogoff, eds., Handbook of international economics (North-Holland, Amsterdam).

Giovannini, A. and M. De Melo, 1993, Government revenue from financial repression, American Economic Review 83, 953–963.

Hall, R.E., 1988, Intertemporal substitution in consumption, Journal of Political Economy 96, 339–357.

Hall, Robert E., 1990, Consumption, In: R. Barro, ed., Studies in business cycle theory (MIT Press: Cambridge, MA).

Hansen, L.P. and R.J. Hodrick, 1980, Forward exchange rates as optimal predictors of future spot rates: An econometric analysis, Journal of Political Economy 88, 829–853.

Heaton, J. and D. Lucas, 1992, Evaluating the effects of incomplete markets on risk sharing and asset pricing, Working paper no. 3491-92-EFA (Alfred P. Sloan School of Management, MIT. Cambridge, MA).

Heaton, J. and D. Lucas, 1995, The importance of investor heterogeneity and financial market imperfections for the behavior of asset prices, Carnegie-Rochester Conference Series on Public Policy 42, 1–32.

Jappelli, T. and M. Pagano, 1989, Consumption and capital market imperfections: An international comparison, American Economic Review 79, 1088–1105.

Jappelli, T., J.-S. Pischke and N.S. Souleles, 1995, Testing for liquidity constraints in Euler equations with complementary data sources, Working paper (University of Pennsylvania, Philadelphia, PA).

Lewis, K.K., 1991, Should the holding period matter in tests of the intertemporal consumption-based CAPM? Journal of Monetary Economics 28, 365–389.

Lewis, K.K., 1995, Puzzles in international finance, In: G. Grossman and K. Rogoff, eds. Handbook of international economics (North-Holland, Amsterdam).

Lewis, K.K., 1996, What can explain the apparent lack of international consumption risk-sharing? Journal of Political Economy 104, 267–297.

Ludvigson, S., 1995, Consumption and credit: A model of time-varying constraints, Working paper (Princeton University, Princeton, NJ).

Newey, W.K. and K.D. West, 1987, A simple, positive semi-definite heteroskedasticity and autocorrelation consistent covariance matrix, Econometrica 55, 703–708.

Obstfeld, M., 1994a, Are industrial-country consumption risks globally diversified? In: L. Leiderman and A. Razin, eds., Capital mobility: The impact on consumption, investment and growth (Cambridge University Press, Cambridge).

Obstfeld, M., 1994b, Risk-taking, global diversification, and growth, American Economic Review 84, 1310–1329.

Obstfeld, M. and K. Rogoff, 1995, The intertemporal approach to the current account, in G. Grossman and K. Rogoff, eds., Handbook of international economics (North Holland, Amsterdam).

Rebelo, S. and C.A. Vegh, 1995, Real effects of exchange-rate-based stabilization: An analysis of competing theories, In: B.S. Bernanke and J.J. Rotemberg, eds., NBER macroeconomics annual, 1995 (MIT Press, Cambridge, MA).

Stockman, A.C. and L.L. Tesar, 1995, Tastes and technology in a two-country model of the business cycle: Explaining international comovements, American Economic Review 85, 168–185.

Summers, R. and A. Heston, 1991, The Penn world table (Mark 5): An expanded set of international comparisons, 1950–1988, Quarterly Journal of Economics 106, 327–368.

Telmer, C.I., 1993, Asset pricing puzzles and incomplete markets, Journal of Finance 48, 1803–1832.

Tesar, L., 1993, International risk-sharing and nontraded goods, Journal of International Economics 35, 69–89.

Tesar, L., 1995, Evaluating the gains from international risk-sharing, Carnegie Rochester Conferences on Public Policy 42, 95–144.

Zeldes, S.P., 1989, Consumption and liquidity constraints: An empirical investigation, Journal of Political Economy 97, 305–346.