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Net Foreign Assets and Imperfect Financial Integration: An Empirical Approach*

by

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Abstract

Researchers have documented extensive empirical evidence on both risk sharing across countries and the uncovered interest rate parity (UIP) condition. This paper involves investigating the empirical implications of imperfectly integrated financial markets resulting from the two phenomena. Under this asset market structure, the net foreign assets (NFA) position of a country affects both the risk-sharing condition and the UIP. Strong evidence exists for Organization for Economic Cooperation and Development (OECD) countries that the NFA contribute to the lack of risk sharing across countries. Similarly, in terms of the UIP, the NFA can capture a time-varying risk premium for a small group of countries over short-term horizons.

Keywords: Net foreign assets, consumption risk sharing, uncovered interest rate parity.

Introduction

Two important issues in international macroeconomics are the apparent lack of risk sharing across countries and the uncovered interest rate parity (UIP) failure. Obstfeld and Rogoff (2000) listed the risk-sharing puzzle among the central unresolved puzzles in international macroeconomics. However, models with integrated financial markets and complete markets result in a risk-sharing condition in which the real exchange rate is as persistent as the ratio of marginal utilities across countries, and the expected change in the nominal exchange rate is proportional to the interest rate differential.

Regarding risk sharing across countries, researchers have studied different data sets and used diverse empirical techniques to test risk-sharing conditions that arise under the assumption of both complete and incomplete markets (Backus & Smith, 1993; Kollmann, 1995; Obstfeld, (1989). Later, Ravn (2001) and Head, Mattina, and Smith (2002) presented evidence that an exogenous incomplete asset market structure (bond economy) does not receive empirical support. Thus, the real exchange rate would not play a role in explaining risk sharing across countries. The evidence results in doubts of the empirical plausibility of theories of international business cycles, which involve assigning a significant role to the real exchange rate in breaking the link of consumption across countries.

Regarding the UIP condition, the hypothesis that interest rate differentials are unbiased predictors of the nominal exchange rate performs poorly. The coefficient on interest rate differentials in exchange rate prediction equations is negative and significant unlike the unitary value that theory predicts. Froot and Thaler (1990), in an extensive empirical testing, offered striking evidence against the UIP. Chinn and Meredith (2005) reported that the UIP does not hold over short horizons but presented evidence suggesting that it may hold over long horizons. Bekaert, Wei, and Xing (2002) suggested that one reason for Chinn and Meredith's (2005) claim that UIP holds better over longer horizons is simply sample choice.

Modern theoretical new open economy macroeconomic (NOEM) models have included assigning an explicit role to the current account and the net foreign assets (NFA) position in the transmission mechanism of shocks across countries, after relegation to a secondary role in previous models.[1] Selaive and Tuesta (2003) examined the role of the NFA in breaking the link between real exchange rate and relative consumption. Their main theoretical contribution was that under the assumption of imperfect financial integration and nontradability, the NFA play a role in explaining the apparent lack of risk sharing across countries.[2] Cavallo and Ghironi (2002), in an overlapping generation model, tried to rationalize the role of the NFA in explaining the permanent U.S. real exchange rate appreciation.

From an empirical perspective, Gagnon (1996) reported a robust long-term relationship between real exchange rate and NFA in a panel of 20 Organization for Economic Cooperation and Development (OECD) economies. Lane and Milesi-Ferretti (2000, 2001a, 2001b) analyzed the determinants of NFA for a large set of economies and offered a variety of theoretical reasons for assuming that some macrovariables should affect NFA. Lane and Milesi-Ferretti also provided evidence of the importance of NFA in determining long-term real interest rate differentials.[3] Bergin (2004) and Rabanal and Tuesta (2006) obtained robust estimates of the debt elasticity parameter in NOEM models by performing structural estimation. Bergin (2004) used maximum likelihood estimation, and Rabanal and Tuesta (2006) performed a Bayesian structural estimation.

The goal of this paper was to investigate the importance of the NFA position in the lack of risk sharing across countries and UIP failure. We tested the imperfect and incomplete asset market structure, closely following Selaive and Tuesta (2003) and Benigno (2001), in which NFA play a crucial role in breaking the link between the real exchange rate and the ratio of marginal utilities and become a debt-elasticity risk premium in the UIP condition. In the benchmark model two risk-free one-period nominal, uncontingent bonds were traded, and the cost of undertaking positions in the international financial markets presented us the opportunity to characterize an imperfect and incomplete asset market structure. Under this asset market structure, assuming deviations from purchasing power parity (PPP), the NFA break the link between the real exchange rate and relative consumptions that characterize models under complete markets. This result arises simultaneously with the NFA directly affecting both the UIP and the risk-sharing relationship. In this context, restrictions in the international financial markets preclude countries from smoothing out consumption, limiting risk-sharing possibilities, and rationalizing the existence of a time-varying risk premium.

In terms of the risk-sharing condition, the findings of the study illustrate that growth factors of consumption and real exchange rates may behave in a manner indicating the significance of the role of the NFA. We found that relatively strong evidence emerged in favor of a risk-sharing relationship with an explicit role for the NFA. For a large sample of countries, the NFA reflect the smooth consumption possibilities bridging the long-lasting gap between the theory and data characterizing previous works. Regarding the UIP relationship, our findings illustrate that the NFA can capture a time-varying risk premium properly only for a small group of countries, showing favorable results in terms of the unbiasedness hypothesis: The interest rate differentials are useful as predictors of short-term movements in exchange rates.

Organization of the remainder of the paper is as follows: The first section contains a brief review of the theoretical approach, which formed the basis of our testing. The following section includes a presentation of some features of the data. Next, the paper includes an empirical discussion of both the risk-sharing and UIP relationships, preceding a discussion of the econometric issues involved in the estimation. The penultimate section contains the results of the study, followed by the conclusion.

A Theory of Imperfect Financial Integration

This section includes a brief presentation of the incomplete asset markets structure (also known as *bond economy*) that Chari, Kehoe, and McGrattan (2002) used in their work. Markets are perfectly integrated under the asset structure of Chari et al. The section also illustrates characterization of an incomplete and imperfect financial assets market structure where the NFA enter explicitly into the risk-sharing relationship and generate deviations from the UIP.

Incomplete Markets

The Standard Approach: Bond Economy

Under both domestic and international complete markets, the ratio of marginal utilities of the two economies equalizes the real exchange rate:[4]

$$q_t = k_o \frac{U_c(C_t^*)}{U_c(C_t)} \quad (1)$$

In Equation 1, k_o is a function of predetermined variables, and $q_t \equiv \frac{S_t P_t^*}{P_t}$ denotes the real exchange rate, with S as the nominal exchange rate, P^* as the foreign price index, and P as the domestic price index. Equation 1 illustrates that the relative consumption across countries is proportional to the real exchange rate.

However, several studies on international business cycles have illustrated an incomplete asset market structure in which the only asset traded internationally is a single, uncontingent nominal bond.[5] Under this asset structure, the risk-sharing condition—see Chari et al. (2002) for further details—reads as follows:

$$E_t\left(\frac{U_c(C_{t+1})}{U_c(C_t)}\frac{P_t}{P_{t+1}}\right) = E_t\left(\frac{U_c(C_{t+1}^*)}{U_c(C_t^*)}\frac{S_t P_t^*}{S_{t+1}P_{t+1}^*}\right) \quad (2)$$

The relation between the real exchange rate and marginal utilities holds in expected first differences.[6] As Equation 2 illustrates, the bond economy allows for a break in the link between real exchange rate and relative consumptions. Although this channel was theoretically promising in addressing the apparent lack of risk sharing, it failed to explain the phenomena.[7] Furthermore, deviations from the UIP were inhibited. On empirical grounds, Obstfeld (1989), Ravn (2001), and Head et al. (2002) cast doubts on the validity of the bond economy approach used by Chari et al. (2002).

Incomplete and Imperfect Financial Integration

This section follows closely the work of Benigno (2001) and Selaive and Tuesta (2003). To break the monotonic relationship between the real exchange rate and relative consumptions, we also generated deviations from the UIP. We assumed that these deviations might stem from the cost of holding foreign bonds, which allowed for the introduction of NFA dynamics into the UIP. Rationalizations of the real exchange rate fluctuations include deviations from the law of one price or the presence of nontraded goods.

The conditions characterizing the allocation of domestic and foreign consumption and holding of nominal bonds appear as follows:[8]

$$U_c(C_t) \quad = \quad (1+i_t)\,\beta \mathrm{E}_t\left\{U_c(C_{t+1})\frac{P_t}{P_{t+1}}\right\} \quad (3)$$

$$U_c(C_t^*) \quad = \quad (1+i_t^*)\,\beta \mathrm{E}_t\left\{U_c(C_{t+1}^*)\frac{P_t^*}{P_{t+1}^*}\right\} \quad (4)$$

$$U_c(C_t) = (1+i_t^*)\,\phi\left(\frac{B_{F,t}S_t}{P_t}\right)\beta \mathrm{E}_t\left\{U_c(C_{t+1})\frac{P_t S_{t+1}}{P_{t+1}S_t}\right\} \quad (5)$$

In the equations, β is the intertemporal discount factor; $\phi(.)$ depends on the real holdings of the foreign assets in the entire economy and, therefore, is taken as given by the domestic household.[9]

Equations 3 and 4 correspond to the Euler equations of the home and foreign countries, respectively. Equation 5 represents household H's Euler equation derived by maximizing the holdings of the nominal bond denominated in foreign currency. From these conditions, we were able to derive the new UIP and the risk-sharing equilibrium condition. The NFA position of the domestic economy affects both the UIP and risk sharing.

Derivation of the UIP involves the difference between the log-linear approximation of Equations 3 and 5, as in the following expression:

$$\widehat{i}_t - \widehat{i}_t^* = E_t(S_{t+1} - S_t) - \delta b_t \quad (6)$$

Notice that Equation 6 incorporates the cost of borrowing in foreign currency and may be consistent with the empirical failure of the UIP.[10] For this study, a time-varying risk premium exists that depends on both the NFA of the domestic economy, b_t, and the cost of bond holdings, δ, which measures the elasticity of the interest rate differential to changes in the NFA position.[11]

The higher the elasticity, the larger the effect of the current account channel on the interest rate differential. The risk premium, δb_t, could be positive or negative depending on the home country being a borrower or a lender in the international assets market. Note that this equation implies a negative relation between the interest rate differential and the NFA of the economy.

We obtained the risk-sharing condition under the imperfect financial integration we imposed by combining the UIP and the corresponding Euler equations for each country:[12]

$$\rho E_t\left(\left(\widehat{C}_{t+1} - \widehat{C}_{t+1}^*\right) - \left(\widehat{C}_t - \widehat{C}_t^*\right)\right) = E_t\left(\widehat{q}_{t+1} - \widehat{q}_t\right) - \delta b_t \quad (7)$$

Equation 7 illustrates the mechanism through which the NFA position affects risk sharing. The characterization of the incomplete asset market structure maintains the gap between relative consumptions that emerge in the bond economy specified in Equation 2, but, in addition, the dynamic of the NFA plays an explicit role.

As long as either asset accumulation or decumulation exists, the NFA will affect the real exchange rate, and the link between the real exchange rate and relative consumptions will be broken. Ceteris paribus, a negative relation exists between the real exchange rate and the NFA; an asset accumulation implies a real exchange rate appreciation. The larger the asset accumulation, the greater the direct effect of the NFA position on the real exchange rate dynamics. Similarly, the larger the cost of undertaking positions in the international financial market, δ, the greater the effect of the NFA on the risk-sharing condition. Finally, if either $\delta \rightarrow 0$ or $b_t=0$ at every period, the risk-sharing relationship amounts to the bond economy.

Features of the Data

Data

All the data collected for the study correspond to quarterly series with the exception of the NFA position, which is only available annually from 1973 to 1998. The series of consumption correspond to quarterly series of private nondurables final consumption at constant prices, obtained from the OECD's Quarterly National Accounts (QNA) and the International Monetary Fund's (IMF) International Financial Statistics (IFS). The series were deflated by the corresponding implicit price deflator for final consumption and multiplied by the nominal exchange rate to express them in terms of U.S. constant dollars.

The database of Lane and Milesi-Ferretti (2001a) provided the series of NFA positions. The NFA were interpolated to obtain quarterly series using the methodology of Chow and Lin (1971).[13] Scaling of the variable occurred by the gross domestic product (GDP) in current dollars for the corresponding year. The quarterly cumulative current account aided in completing the data for the period 1999 to 2001.

The definition of bilateral real exchange rates was the nominal exchange rate multiplied by the ratio of foreign to domestic prices. The IFS provided the nominal exchange rate data. Prices were defined as the implicit deflators for the consumption variables. The IFS provided the real effective exchange rate for the period January 1975 to April 2001. To complete the sample, with a starting point of 1973, we defined the real effective exchange rate as the nominal effective exchange rate multiplied by the ratio of the aggregate OECD prices to domestic prices. The nominal effective exchange rates appeared in the IFS data. The interest rate, obtained from the database of the Bank of International Settlements, corresponded to 3-month and 12-month euro-currency yields expressed in annual terms.

Description of the Data

Under the standard assumption of separability in the utility function and assuming that purchasing power parity (PPP) holds, a complete asset market assumption will imply a perfect cross-correlation of consumptions across countries. Table 1 illustrates the cross-country correlations of consumption growth rates. For most of the countries, the correlation is very low and increases when calculated with respect to an aggregate of OECD consumption, consistent with Ravn's (2001) findings.

As indicated in the previous section, the real exchange rate introduces a wedge between the relative consumptions across countries; thus, consumption correlations do not necessarily have to be perfect. The bond economy will predict a positive relation between the fluctuations in the relative consumption growth rates and those of the real exchange rate. Table 2 depicts the cross-correlations between bilateral (and effective) real exchange rates and consumption growth rates for 12 OECD economies for the period January 1970 to April 2001.[14] The cross-corre-

Table 1
Cross-country consumption correlation

	AUS	CAN	FIN	FRN	ITY	JPN	NWY	SPN	SWT	UK	US	OECD
AUL	0.11	0.14	0.37	0.05	0.15	0.04	-0.08	0.33	0.32	0.14	-0.03	0.15
AUS		0.02	0.00	0.00	0.20	0.05	0.04	0.33	0.21	0.04	0.02	0.17
CAN			0.06	0.11	0.28	0.06	-0.11	0.29	0.11	0.26	0.36	0.31
FIN				-0.24	0.12	0.08	0.17	0.33	0.34	0.54	0.14	0.34
FRN					0.00	0.23	-0.68	0.01	-0.17	0.07	0.13	0.08
ITY						0.03	0.06	0.51	0.36	0.15	-0.10	0.32
JPN							-0.17	0.10	0.05	0.17	0.17	0.22
NWY								0.04	-0.01	-0.05	0.02	-0.09
SPN									0.30	0.26	0.11	0.44
SWT										0.09	-0.07	0.22
UK											0.36	0.47
US												0.28

Notes: Numbers correspond to Corr($C^i_{t+1} - C^i_t$, $C^j_{t+1} - C^j_t$).

Table 2
Unconditional Correlation: Risk Sharing

	Change in Real Effective Exchange Rate with#:		Change Real Bilateral Exchange Rate with#:	
Country	$\Delta C^{R\text{-}ROW}_{t+1}$	b_t	$\Delta C^{R\text{-}US}_{t+1}$	b_t
Australia	-0.02	-0.04	0.01	0.07
Austria	-0.18	-0.05	-0.23	-0.11
Canada	-0.11	0.04	-0.20	0.00
Finland	0.02	0.00	0.30	0.05
France	-0.02	-0.06	0.51	0.01
Italy	0.04	-0.06	0.32	-0.07
Japan	0.18	-0.04	0.16	0.04
Norway	0.01	-0.14	0.35	0.09
Spain	-0.19	-0.14	0.44	-0.16
Switzerland	-0.06	-0.16	0.35	-0.22
UK	0.03	0.12	-0.19	-0.15
US	-0.10	0.01		

Notes: $\Delta C^{R\text{-}ROW}$ is the relative consumption growth rate relative to the Rest of the World; $\Delta C^{R\text{-}US}$ is the relative consumption growth rate relative to the U.S; and b corresponds to NFA/GDP. paper. # corresponds to the change in the real exchange rate defined in the data section (q_t- q_{t-1}).

lations are quite low and negative in most of the cases, so the data may not support the bond economy. This finding confirms the econometric estimations by Ravn (2001) and Head et al. (2002).[15]

The theory tested in this paper includes considering the NFA as a key determinant of the lack of risk sharing across countries (see Equation 7). Figures 1 and 2 depict, for the period 1973 to 2001, the NFA position in relation to the real effective and bilateral exchange rate, respectively. The bold line represents the real exchange rate, and the dotted line indicates NFA.

Most of the real exchange rate series exhibit large swings around a slowly drifting mean. The NFA drift upward for Japan, Norway, and Switzerland and downward for Australia and the United States, with little trend in the remaining countries. The theory we tested would predict a positive correlation between the expected growth rate of the real exchange rate and NFA.

Table 2 shows the correlations between the growth rate of the real exchange rate and the NFA for the set of countries. Mixed results included positive and negative correlations, most of them quite low. The results provide preliminary evidence in favor of a theory in which NFA may play a role for some economies.

We performed a similar exercise for the UIP condition. Table 3 indicates the cross-correlations of both interest rate differentials and NFA positions with the change in the nominal exchange rate. The correlations are negative for most of the countries, using a short-term interest rate differential, although they increase when using the

Table 3
Unconditional Correlation: UIP

	Change in Nominal Exchange Rate with#:		
Country	$\Delta i_t^{R\text{-}3m}$	$\Delta i_t^{R\text{-}12m}$	b_t
Australia	0.06	-0.53	-0.10
Austria	-0.12	0.14	-0.11
Canada	-0.10	-0.18	0.03
Finland	0.33	0.15	0.06
France	-0.11	0.36	0.05
Germany	-0.12	0.26	-0.10
Italy	0.10	0.38	-0.05
Japan	-0.42	-0.06	0.09
Netherlands	-0.17	0.10	-0.08
Norway	-0.16	-0.09	0.05
Spain	0.08	0.27	-0.11
Switzerland	-0.18	-0.01	-0.14
Sweden	-0.09	-0.06	0.16
UK	-0.19	0.05	-0.15

Notes: $\Delta i_t^{R\text{-}3m}$ and $\Delta i_t^{R\text{-}12m}$ are the 3- and 12-months interest rate differentials in euro currency yields (see data section for details on this variable); and b corresponds to *NFA/GDP*. # corresponds to the change in the nominal exchange rate defined in the data section (S_t- S_{t-1}). The change in the nominal exchange rate corresponds to the same maturity of the interest rate differential.

12-month interest rate differential.[16] However, the NFA position correlates positively with the expected change in the exchange rate for 6 out of 14 countries.

The evidence indicates that the risk-sharing hypothesis and the UIP condition may assign a role to the NFA position empirically. However, the correlation analysis does not constitute robust empirical testing. Hence, in the next section we evaluate empirically the role of the NFA in explaining both risk sharing and the UIP failure.

Empirical Discussion

Obstfeld (1989) first derived and tested a risk-sharing hypothesis in a setup where PPP did not hold. Ravn (2001), following a similar approach, assumed that countries can borrow and lend freely at the same nominal interest rate. Both authors did not find evidence in favor of the bond economy setup.[17] Head et al. (2002) extended the research by testing utility functionals with stochastic discount rates that are consistent with a stationary distribution of wealth when markets are exogenously incomplete and augmented utility with external habit persistence as applied by Campbell and Cochrane (1999). Their empirical findings did not support any of these extensions.

The basis of the econometric approach for the study was the work of Kollmann (1995), who studied the relation between consumption and the real exchange rate using the generalized method of moments estimation procedure (GMM).[18] However, our approach differed from Kollmann's work in several ways. First, our study involved investigating a broader data set. Second, the research included testing a different risk-sharing condition, which the NFA enter explicitly. Third, the study involved testing the UIP condition that arises from the asset market structure suggested previously.

The current account may indeed occupy an important role in the international transmission mechanism of shocks. Gagnon (1996) presented evidence of a significant and robust long-term relationship between the real exchange rate and the NFA. In addition, Lane and Milesi-Ferretti (2000) argued that NFA have a strong impact on the relative price of nontraded goods and, therefore, on the real exchange rate dynamics.[19]

With respect to the UIP, extensive empirical studies have resulted in the rejection of the hypothesis that interest rate differentials are unbiased predictors of future exchange rate movements. The UIP predicts that high yield currencies should be expected to depreciate, and, ceteris paribus, a real interest rate increase should appreciate the currency. When inefficient markets or short-term market frictions prevent an immediate, complete response of the exchange rate to an interest rate change, short-term deviations of UIP may occur while long-horizon UIP holds.

Chinn and Meredith (2005) tested the UIP hypothesis on longer maturity bonds for the United States, Germany, Japan, and Canada and found evidence that the longer the maturity, the better the interest rate differential explains the future exchange rate variations.[20] Chinn and Meredith interpreted this as meaning that any risk premium is very stable over long horizons. However, Bekaert et al. (2002) posited that one reason for Chinn and Meredith's (2005) claim that UIP holds better at longer horizons is simply sample choice. Finally, Lane and Milesi-Ferretti (2001b) discovered a strong long-term link between the real interest rate differential and NFA for a large sample of countries.

Econometric Issues

Persistence of NFA

The NFA position is a variable that exhibits high persistence in the data. The largest autocorrelation root for most of the countries in our sample is in the interval [0.9,0.98] (not reported to save space). This is not an isolated characteristic of the NFA position and is observed for a wide set of macroeconomic variables (Stock & Watson, 1996).

However, the study's sample period started with January 1973, which limited the number of observations to no more than 100. In this context, to rely on tests of the null hypothesis of a stationary process, such as the KPSS and LMC tests, that are based on conventional asymptotic critical values, may be misleading and result in rejection of the null hypothesis. Conversely, Caner and Kilian (2001) have shown that one can overcome the problem with tests that rely on the null hypothesis of unit root when one corrects the critical values for finite-sample or bootstrap critical values.

Thus, we employed the efficient DF-GLS test (Elliott, Rothenberg, & Stock, 1996) of the unit root null hypothesis using finite-sample critical values. Creating finite-sample critical values to compare with the statistic generated by the test involved following the work of Caner and Kilian (2001) closely. Table 4 illustrates the results.

After applying the test to the NFA for the set of countries in our sample, we rejected the unit root null at 10% for 7 out of 14 countries. The result indicates that the NFA are not only theoretically but also empirically mean-reverting for some countries. Therefore, the countries that cannot reject the null of unit root are strongly influenced by a small sample-size problem. [21]

Estimation Procedure

We used the GMM procedure developed by Hansen (1982). Under this estimation procedure, we minimized a criterion function derived by imposing at least as many moment conditions as parameters to be estimated. If instruments correlate poorly with the endogenous variables, they provide limited ability to discriminate among various parameter values, so GMM inferences are misleading, and a weak identification result could arise. According to increasing the number of instruments will not solve the problem.

Despite the evolving nature of the literature, some useful methods emerge to address concerns about weak identification. Stock and Wright (2000) developed a test to check for the presence of weak instruments. Following the example of the authors, we implemented a fully robust test for weak identification (*S*-set test*)*. We used the robust continuous-updating GMM estimator, which minimizes the following objective function:

$$S(\Psi) = \left[\frac{1}{\sqrt{T}}\sum_{t=1}^{T}\phi_t(\Psi)\right]' V(\Psi)^{-1}\left[\frac{1}{\sqrt{T}}\sum_{t=1}^{T}\phi_t(\Psi)\right]$$

In the equation, Ψ is the set of parameters to be estimated; $\phi_t(\Psi)=h(Y_t,\Psi)\otimes Z_t$, where $h(Y_t,\Psi)$ is the orthogonality condition, and Z_t is a vector of instruments; $V(\Psi)$ is the robust variance-covariance matrix.

We juxtaposed the conventional 90% confidence ellipse with the 90% *S*-set. The *S*-set contains all parameters that pass the 90% χ^2_k test, constructed according to Theorem 2 in Stock and Wright (2000). Loosely speaking, if the *S*-set were contained in the 90% confidence ellipse, the instruments would not be weak.

Results

Testing a Risk-Sharing Condition Under Imperfect Financial Integration

Our incomplete and imperfect asset market structure delivered the following orthogonality condition for estimation:[22]

$$E_t\left\{\rho\Delta\widehat{C}^R_{t+1} - \Delta\widehat{q}_{t+1} + \delta b_t\right\}\otimes Z_t = 0 \tag{8}$$

In the equation, Z_t corresponds to the vector of instruments, $\Delta\widehat{C}^R_{t+1}$ is the growth rate of relative consumptions, and $\Delta\widehat{q}_{t+1}$ is the growth rate of the real exchange rate. Finally, b_t stands for the ratio of NFA in current dollars scaled by the GDP in current dollars.

We examined the risk-sharing condition, Equation 8, using quarterly data for a set of 12 countries and performed three estimations. First, we considered the real effective exchange rate and relative consumption growth rates with respect to the rest of the world (RoW). Second, we examined the same risk sharing on a country-pair basis with respect to the United States. Finally, we performed a balanced panel for both cases.

Table 5 shows the results of the estimation country-RoW. By way of contrast, the second and third columns display the estimation that corresponds to the bond economy. The estimated coefficient of risk aversion, ρ, is negative for 7 of the 12 countries and is positive and significant for only 2 countries (Japan and Italy). The associated p values of the J statistics were above 0.10 in all cases, so

Table 4
Efficient DF-GLS Test for the NFA

Country	*DF-GLS*	Reject I(1) null (5 or 10%) #	Sample Period
Australia	-1.732	*yes*	1970:1 – 2001:4
Austria	-4.562	*yes*	1970:1 – 2001:4
Canada	-1.693	*yes*	1971:4 – 1997:3
Finland	-1.879	*yes*	1970:1 – 2001:4
France	-1.715	*yes*	1970:1 – 2000:4
Germany	-0.516	*no*	1970:1 – 2001:4
Italy	-0.437	*no*	1970:1 – 2001:4
Japan	0.002	no	1970:1 – 2001:4
Netherlands	-0.433	*no*	1970:1 – 2001:4
Norway	0.354	*no*	1970:1 – 2001:4
Spain	-1.881	*yes*	1970:1 – 2000:4
Sweden	-0.165	*no*	1980:4 – 2001.4
Switzerland	-0.948	*no*	1970:1 – 2001:4
UK	-1.801	*yes*	1970:1 – 2001:4

Notes: - Sample period: 1970:1-2001:4. We have allowed for 8 lags to construct the statistic, and the finite sample critical values.
(#) Finite-sample critical values -2.09[-1.69] at 5[10]%.

we did not reject the null of overidentifying restrictions. The results indicate at least weak evidence in favor of the bond economy.

Next, we tested the proposed risk-sharing relationship, Equation 8. The results appear in the last columns of Table 5. The first striking result is that the estimate of the risk-aversion parameter is positive and significant for 7 out of 12 countries, which may suggest that the instruments associated with the NFA positions are helping to identify the risk-aversion parameter and to capture some aspects of smooth consumption possibilities.[23] The second important result is the positive and significant value of the cost-of-bond-holding parameter, δ, for 5 countries in the sample.[24] The associated p values of the J statistics were all above 0.10. Finally, for 3 countries (Australia, Japan, and Norway), both parameters, ρ and δ, are positive and significant.

We examined the possibility that the results may be driven by weak identification problems by constructing the conventional 90% confidence ellipse with the 90% S-set described previously. The last column of Table 5 illustrates the results, and the S-set test results appear in Figure 3.[25] Under the reasonable assumption that the risk-aversion parameter is not too large, as indicated in previous empirical evidence (Yogo, 2004), our estimations may not be driven by weak identification.

Table 5

Risk-Sharing with the Rest of the World

$$\rho\,(C_{t+1} - C_t - (C^*_{t+1} - C^*_t)) = q_{t+1} - q_t \qquad (1)$$

$$\rho(C_{t+1} - C_t - (C^*_{t+1} - C^*_t)) = q_{t+1} - q_t - \delta b_t \qquad (2)$$

Country	(2) ρ	(2) J-stat	(3) ρ	(3) δ	(3) J-stat	(3) S-Set
			Country-RoW			
Australia	-0.198	0.80	2.504***	0.006**	0.80	No empty
	(1.247)		*(0.621)*	*(0.003)*		
Austria	0.234	0.15	1.204*	-0.022**	0.72	Empty
	(0.556)		*(0.494)*	*(0.005)*		
Canada	1.449	0.94	1.328*	-0.001	0.91	Empty
	(1.443)		*(0.775)*	*(0.004)*		
Finland	-1.308*	0.54	-2.005***	-0.008	0.61	Empty
	(0.767)		*(0.548)*	*(0.004)*		
France	-0.384	0.85	-0.746***	0.071***	0.78	Empty
	(0.570)		*(0.257)*	*(0.023)*		
Italy	0.883**	0.50	0.499**	0.025	0.71	No empty
	(0.392)		*(0.294)*	*(0.020)*		
Japan	2.643***	0.44	2.995***	0.018*	0.98	No empty
	(1.157)		*(0.655)*	*(0.011)*		
Norway	-0.101	0.54	0.335**	0.007**	0.85	No empty
	(0.450)		*(0.168)*	*(0.003)*		
Spain	-2.17***	0.66	-2.061**	0.012*	0.60	No empty
	(0.642)		*(0.441)*	*(0.006)*		
Switzerland	0.914	0.51	1.822***	-0.003	0.74	Empty
	(0.568)		*(0.441)*	*(0.004)*		
U.Kingdom	-1.589	0.64	-1.412	0.017	0.14	No empty
	(1.470)		*(1.195)*	*(0.016)*		
United States	-0.276	0.95	-2.641***	-0.014*	0.93	No empty
	(1.022)		*(0.571)*	*(0.007)*		
Panel 1/	0.335**	0.35	0.623***	0.001**	0.97	
	(0.133)		*(0.023)*	*(0.000)*		

Notes: Estimations by *GMM*, Standard Errors are reported in parenthesis modified by Newey-West correction. Instruments are lagged relative consumption growth rate, lagged real exchange growth rate and lags of net foreign asset position. J-Statistic is the significance level of a test of the overidentifying restrictions. *S-set* tests for weak instruments: "empty" set implies weak identification. Sample period: 1970:4-2000:4. *, (**), [***] Significance at 10%, (5%), [1%].

1/ All countries but USA and Australia. The sample runs from 1975:2 2000:4.

The last row of Table 5 reflects the panel GMM estimation. Both the coefficient of risk-aversion and the cost-of-bond-holding parameters are significant and have the correct sign. Thus, the results support a theory of imperfectly integrated financial markets.

Finally, Table 6 indicates the results of the estimation on a country-by-country basis. Again, the second and third columns reflect the bond economy. The coefficient of risk aversion is significant and has the correct sign only for 3 out of 11 countries. For the other 8 countries, the parameter is either not significant or negative. Including the NFA in the equilibrium condition improved the estimations considerably. The estimate of the risk-aversion parameter is positive and significant for 8 out of 11 countries, and the estimate of the cost of bond holdings is positive and significant for 7 countries.[26]

The associated p values of the J statistics were above 0.10 in all cases.[27] In most of the cases, 8 out of 11, the results were not driven by weak identification problems (see Figure 4). Again, the panel results support the theory because the estimates of parameters ρ and δ are positive and significant. The results indicate that a theory of imperfect financial integration may work better on a country-by-country basis rather than on a country-RoW basis.

Table 6
Bilateral Risk-Sharing

$$\rho(C_{t+1} - C_t - (C^*_{t+1} - C^*_t)) = q_{t+1} - q_t \qquad (3)$$
$$\rho(C_{t+1} - C_t - (C^*_{t+1} - C^*_t)) = q_{t+1} - q_t - \delta b_t \qquad (4)$$

Country	(3) ρ	(3) J-stat	(4) ρ	(4) δ	(4) J-stat	(4) S-Set
			Country-USA			
Australia	-0.273	0.25	1.583***	0.004*	0.95	No empty
	(0.781)		*(0.465)*	*(0.003)*		
Austria	0.799	0.74	1.939*	-0.031	0.77	No empty
	(0.556)		*(1.022)*	*(0.021)*		
Canada	-2.212**	0.29	-0.496	0.004**	0.81	No empty
	(1.085)		*(0.394)*	*(0.002)*		
Finland	0.946**	0.30	1.264***	-0.007	0.70	No empty
	(0.365)		*(0.372)*	*(0.007)*		
France	-0.040	0.97	0.484**	0.089*	0.95	No empty
	(0.493)		*(0.210)*	*(0.053)*		
Italy	0.212	0.78	0.312**	0.060*	0.88	No empty
	(0.237)		*(0.154)*	*(0.032)*		
Japan	2.320**	0.40	3.727***	0.044***	0.28	Empty
	(0.933)		*(0.933)*	*(0.017)*		
Norway	0.106	0.64	-0.076	0.015**	0.88	No empty
	(0.332)		*(0.235)*	*(0.006)*		
Spain	1.041***	0.46	1.118***	-0.038***	0.65	Empty
	(0.315)		*(0.225)*	*(0.013)*		
Switzerland	0.092	0.79	1.004***	0.010*	0.60	Empty
	(0.293)		*(0.146)*	*(0.005)*		
U. Kingdom	-0.484	0.49	0.636	-0.039**	0.60	No empty
	(1.522)		*(0.670)*	*(0.019)*		
Panel 1/	0.119	0.29	0.529**	0.003**	0.77	
	(0.116)		*(0.046)*	*(0.001)*		

Notes: Estimations by *GMM*, Standard Errors are reported in parenthesis modified by Newey-West correction. Instruments are lagged relative consumption growth rate, lagged real exchange growth rate and lags of net foreign asset position. J-Statistic is the significance level of a test of the overidentifying restrictions. *S-set* tests for weak instruments: “empty” set implies weak identification. Sample period: 1970:4-2000:4. *, (**), [***] Significance at 10%, (5%), [1%].
1/ The sample runs 1975:02 2000:04.

Overall, the tested risk-sharing condition worked well. We have highlighted the importance of the NFA in explaining the lack of risk sharing across countries, and in general, the structural estimates for more than half of the countries in the sample are in line with the predictions of the theory. Growth factors of consumption and real exchange rates appear to behave in a manner that may be consistent with the assumptions implicit in the incomplete and imperfect market structure.

Testing the UIP

We examined the role of the NFA in explaining the UIP condition in the short term rather than testing the UIP at different horizons.[28] The market structure outlined previously, in a regression context, delivers the following orthogonality condition:

$$E_t\left\{\alpha+\beta\left(\widehat{i}_t-\widehat{i}_t^*\right)-E_t\Delta S_{t+1}+\delta b_t\right\}\otimes Z_t=0 \quad (9)$$

In the equation, Z_t corresponds to the vector of instruments, ΔS_{t+1} is the growth rate of the bilateral nominal exchange rate, and $\left(\widehat{i}_t-\widehat{i}_t^*\right)$ is the interest rate differential.

Table 7

Uncovered Interest Parity: Maturity 3 months

$$\Delta S_{t,t+k} = \alpha + \beta(i_{t,k} - i^*_{t,k}) + \varepsilon_{t,k} \quad (5)$$
$$\Delta S_{t,t+k} = \alpha + \beta(i_{t,k} - i^*_{t,k}) + \delta b_{t,k} + \varepsilon_{t,k} \quad (6)$$

Country	(4) Estimate β	(5) *Maturity*: 3 months Estimate β	Estimate δ
Australia	-0.302***	-0.371***	-0.115
	(0.457)	*(0.452)*	*(0.126)*
Austria	-1.047***	-1.286	0.287
	(0.667)	*(0.876)*	*(0.591)*
Canada	-0.706***	-1.236***	0.033
	(0.289)	*(0.312)*	*(0.109)*
Finland	-1.047***	2.189	1.189†††
	(1.028)	*(1.243)*	*(0.236)*
France	-0.191	1.346	2.471†††
	(0.957)	*(1.109)*	*(0.993)*
Germany	-0.865***	0.102	-1.176††
	(0.944)	*(0.848)*	*(0.471)*
Italy	0.856	2.543**	2.081†††
	(0.662)	*(0.733)*	*(0.584)*
Japan	-5.777***	-4.198***	0.558†††
	(0.935)	*(0.730)*	*(0.243)*
Netherlands	-1.728***	-1.143***	-0.333
	(0.841)	*(0.650)*	*(0.376)*
Norway	-0.982***	-1.024***	-0.023
	(0.621)	*(0.633)*	*(0.076)*
Spain	0.817	1.067	-0.337
	(0.553)	*(0.642)*	*(0.457)*
Sweden	-2.831***	-2.822***	0.587†††
	(0.923)	*(0.926)*	*(0.262)*
Switzerland	-1.587***	-0.875**	-0.355
	(0.686)	*(1.019)*	*(0.374)*
UK	-2.090***	-1.259**	-0.230
	(1.013)	*(0.941)*	*(0.271)*

Notes: - Estimations by *GMM*, Standard Errors in parenthesis modified by Newey-West correction. All *p-values* of *J- statistics* are above 0.1. *(**)[***]. Different from null of unity at 10%(5%)[1%]. †(††)[†††]. Different from null of zero at 10%(5%)[1%]. Bilateral Nominal Exchange Rate in terms of US dollars. Interest Rate differential in Eurocurrency yields. *NFA* country-RoW. Sample period: 1980:1-2001:4.

We estimated the UIP, Equation 9, by GMM for two different horizons with 3- and 12-month maturity bonds (Chinn & Meredith, 2005). At this stage, we did not claim that the NFA may help to predict exchange rate movements but wanted to assess whether a significant role exists for the NFA in explaining exchange rate movements as our theory would predict.[31] In particular, we wanted to assess the significance and sign of the parameter δ in Equation 9.

Using constant-maturity 3-month yields for 14 countries, we implemented regressions of the form of Equation 9 from January 1980 to April 2001. By way of contrast, we also displayed the estimates of the UIP condition analyzed in previous studies and used the results obtained by Chinn and Meredith (2005) as a benchmark. Chinn and Meredith estimated the UIP for short-term horizons, 3 months and 12 months, for the deutschmark, Japanese yen, UK pound, French franc, Italian lira, and Canadian dollar. In this study, we used exactly the same data set as Chinn and Meredith. The exchange rates of each country were expressed in terms of the U.S. dollar, and the 3- and 12-month movements in the exchange rate were regressed against differential in euro currency yields of the corresponding maturity. Because 12-horizon data at a quarterly frequency may have led to MA(3) in the residuals, we used the Newey-West correction to obtain robust standard errors.

Table 8

Uncovered Interest Parity: Maturity 12 months

$$\Delta S_{t,t+k} = \alpha + \beta\,(i_{t,k} - i^{*}_{t,k}) + \varepsilon_{t,k} \qquad (7)$$

$$\Delta S_{t,t+k} = \alpha + \beta\,(i_{t,k} - i^{*}_{t,k}) + \delta b_{t,k} + \varepsilon_{t,k} \qquad (8)$$

Country	(4) *Estimate β*	(5) *Maturity*: 12 months *Estimate β*	*Estimate δ*
Australia	-2.306***	-3.667***	-0.592†††
	(0.650)	*(0.656)*	*(0.222)*
Austria	0.342	1.715	-1.214†††
	(0.955)	*(0.856)*	*(0.273)*
Canada	0.085***	0.001***	0.094
	(0.145)	*(0.152)*	*(0.113)*
Finland	1.071	2.378	0.972†††
	(1.067)	*(1.016)*	*(0.377)*
France	1.675	1.152	1.841†
	(0.396)	*(0.379)*	*(1.157)*
Germany	-0.776***	0.082	-1.617†††
	(0.811)	*(0.650)*	*(0.353)*
Italy	0.856	1.694	1.278†
	(0.662)	*(0.263)*	*(0.840)*
Japan	-0.276***	2.778***	0.775†††
	(0.600)	*(1.254)*	*(0.273)*
Netherlands	-1.353***	-0.972	1.741
	(0.787)	*(1.995)*	*(2.529)*
Norway	0.420	0.422	-0.021
	(0.512)	*(0.725)*	*(0.160)*
Spain	1.682	2.612**	-1.303††
	(0.523)	*(0.697)*	*(0.622)*
Sweden	-0.864***	-0.826***	0.626††
	(0.586)	*(0.480)*	*(0.315)*
Switzerland	-0.948***	-1.086**	-0.415††
	(0.872)	*(0.810)*	*(0.193)*
UK	-1.043***	1.133	-0.699†††
	(0.560)	*(0.580)*	*(0.271)*

Notes: Estimations by *GMM*, Standard Errors in parenthesis modified by Newey-West correction. All *p-values* of *J- statistics* are above 0.1. *(**)[***] different from null of unity at 10%(5%)[1%]. †(††)[†††] different from null of zero at 10%(5%)[1%].
Bilateral Nominal Exchange Rate in terms of US dollars. Interest Rate differential in Eurocurrency yields. *NFA* country-RoW. Sample period: 1980:1-2001:4.

Table 7 indicates the estimations for 3-month maturity. The first column shows that the UIP condition is rejected in most of the countries when the NFA position is not included as a regressor, which is in line with the results reported by Chinn and Meredith (2005). The results illustrate the failure of the UIP and confirm the results of Froot and Thaler (1990). If UIP holds, the slope coefficient should not be statistically different from one. Regarding the constant term, nonzero values may be explained by Jensen's inequality but are not presented to save space. The estimate slope coefficient, β, has a negative sign for 12 out of 14 countries. Only France and Italy present coefficients that are not statistically different from one. The second and third columns of Table 7 present the estimate slope and risk-premium coefficients of the UIP condition, Equation 9.

The first remarkable result was that the parameter δ is positive and significant for 5 countries. Moreover, for 3 economies (Finland, France, and Italy), both the slope and the risk-premium coefficients have the right sign and are significant at the 10% level. The findings contrast with the results of Chinn and Meredith (2005), who reported the failure of the UIP for most of the currencies analyzed at 3-month horizons. For Sweden and Japan, the risk-premium coefficient is positive and significant, although the slope coefficient has the wrong sign. For the rest of the countries, the slope coefficient moves in the right direction, but the risk-premium parameter has the wrong sign and/or is positive but not significant.

We performed the same estimation for longer maturity bonds, and the results at 12-month maturity appear in Table 8. Again, some support exists for our theory. The slope parameter is significant and has the correct sign for 7 countries. The results improved slightly with respect to the 3-month horizon. The risk-premium parameter is also positive for 7 countries and is significant for 5.

In summary, NFA may be useful predictors of short-term movements in exchange rate for some countries. They may explain the observed variance in exchange rates for the period analyzed. However, the findings indicate that NFA may not be an appropriate measure of time-varying risk premium for an important subsample of countries.

Joint Test of Risk Sharing and the UIP

We performed a tighter test of the imperfect and incomplete asset market structure presented previously. We implemented a joint GMM estimation of the risk-sharing and UIP equilibrium conditions. Under the joint estimation, we made the risk-premium parameter, δ, the same in both equations. One of the limitations of this approach was that the sample size was limited to the sample period used in the UIP estimation while we increased the number of moment conditions and parameters to be estimated. To perform a balanced estimation, we restricted the number of countries in the sample to 10. The results appear in Table 9.

The estimated risk-aversion parameter, ρ, is positive and significant for 7 countries. The slope coefficient of the UIP condition has the right sign for 4 countries. Finally, parameter δ, intended to capture the time-varying risk premium generated by the NFA, is significant at 10% for 5 economies.

The results of the estimations indicate two main ideas. First, a significant role exists for the NFA position in explaining the lack of risk sharing across countries. Second, a significant link appears to exist between the UIP puzzle and the apparent lack of risk sharing across countries.

Concluding Remarks

This study involved examining the empirical implications of incomplete asset markets and imperfect financial integration in explaining both the apparent lack of risk sharing across countries and the UIP failure. Researchers have documented the empirical failure of most of the theoretical models under the assumption of perfect integration, even when allowing for both exogenously and endogenously incomplete markets, habit persistence, and different forms of utility functions. Evidence of the importance of NFA in explaining the transmission of shocks across countries prompted us to consider the implications of imperfect financial integration in international macroeconomics.

The results of the paper contrast with results of previous studies based on the assumption of complete markets. First, in the study, evidence emerged that growth factors of consumption and real exchange rates may behave in a manner consistent with a significant role for NFA for a large sample of OECD countries. In this sense, NFA are a key element in explaining the apparent lack of risk sharing across countries. Second, for a small group of countries, NFA capture a time-varying risk premium and yield a positive slope coefficient for the interest rate differential at short-term horizons, a result that contrasts with the findings of Chinn and Meredith (2005). In this sense, the interest rate differential could be a useful predictor of short-term movements in the nominal exchange rate when accompanied by the NFA position.

In conclusion, considering a theory where the NFA position affects both the risk sharing across countries and the UIP condition seems reasonable. The findings illustrate that because the NFA help to explain nominal exchange rate movements, an important avenue to investigate is the predictability power of NFA following the germinal contribution of Meese and Rogoff (1983). In this line, to overcome the high persistence of the NFA in the empirical testing by a suitable transformation may be a good alternative.

Table 9

Joint Estimation of the UIP and Risk-Sharing Condition.

$$\rho[C_{t+1} - C_t - (C^*_{t+1} - C^*_t)] - (q_{t+1} - q_t) + \delta b_t = 0$$
$$\alpha + \beta(i_{t,k} - i^*_{t,k}) + \delta b_{t,k} - \Delta S_{t,t+k} = 0$$

	Joint Estimation		
Country	**Estimate ρ**	**Estimate β**	**Estimate δ**
Australia	-0.254	0.445	-0.013†††
	(1.328)	*(0.401)*	*(0.004)*
Austria	3.979†	-0.742**	-0.074†††
	(2.323)	*(0.273)*	*(0.031)*
Canada	1.164†††	-0.758***	-0.007†
	(0.152)	*(0.214)*	*(0.004)*
France	-0.738†	-1.019***	0.140††
	(0.395)	*(0.670)*	*(0.064)*
Italy	1.141†††	0.746	0.072†
	(0.271)	*(0.545)*	*(0.048)*
Japan	2.238†††	-4.426***	0.079†††
	(0.454)	*(0.758)*	*(0.011)*
Norway	0.094	1.033	0.011†
	(0.179)	*(0.536)*	*(0.006)*
Spain	0.898††	1.659	-0.039††
	(0.431)	*(0.423)*	*(0.016)*
Switzerland	2.194†††	-1.955***	0.025†††
	(0.333)	*(0.342)*	*(0.005)*
UK	-1.821	-1.336**	-0.009
	(1.219)	*(0.833)*	*(0.020)*

Notes: Estimations by *GMM*, Standard Errors in parenthesis modified by Newey-West correction. *(**)[***] different from null of unity at 10%(5%)[1%]. †(††)[†††] different from null of zero at 10%(5%)[1%].

Bilateral Nominal Exchange Rate in terms of US dollars. 3-months maturity Interest Rate differential in Eurocurrency yields. *NFA* country-RoW. Balance Sample period: 1980:1-2001:4.

References

Backus, D., & Smith, G. (1993). Consumption and real exchange rates in dynamic economies with non-traded goods. *Journal of International Economics, 35*(3-4), 297-316.

Bekaert, G., Wei, M., & Xing, Y. (2002). Uncovered interest parity and the term structure. *Journal of International Money and Finance, 26*(6), 1038-1069.

Benigno, P. (2001). *Price stability with imperfect financial integration*. Unpublished manuscript, New York University.

Bergin, P. (2004). *How well can the new open economy macroeconomics explain the exchange rate and the current account?* Unpublished manuscript, University of California at Davis.

Campbell, J., & Cochrane, J. H. (1999). By force of habit: A consumption-based explanation of aggregate stock market behavior. *Journal of Political Economy, 107,* 205-251.

Caner, M., & Kilian, L. (2001). Size distortions of tests of the null hypothesis of stationarity: Evidence and implications for the PPP debate. J*ournal of International Money and Finance, 20*(5), 639-657.

Cavallo, M., & Ghironi, F. (2002). Net foreign assets and the exchange rate: Redux revived. *Journal of Monetary Economics, 49*(5), 1057-1097.

Chari, V. V., Kehoe, P., & McGrattan, E. (2002). Can sticky prices explain the persistence and volatility of real exchange rate? *Review of Economic Studies, 69*(3), 533-563.

Chinn, M., & Meredith, G. (January 2005). Testing uncovered interest parity at short and long horizons during the Post-Bretton Woods Era (NBER Working Paper No. 11077). National Bureau of Economic Research.

Chow, G., & Lin, A. (1971). Best linear unbiased distribution and extrapolation of economic time series by related series. *Review of Economics and Statistics, 53*(4), 372-375.

Elliott, G., Rothenberg, T., & Stock, J. (1996). Efficient tests for an autoregressive unit root. *Econometrica, 64*(4), 813-836.

Froot, K., & Thaler, R. (1990). Foreign exchange. *Journal of Economic Perspectives, 4*(3), 179-192.

Gagnon, J. (1996). Net foreign assets and equilibrium exchange rates: Panel evidence (International Finance Discussion Paper No. 574). Washington, D.C.: Board of Governors of the Federal Reserve System.

Hansen, L. (1982). Large sample properties of generalized method of moments estimators. *Econometrica*, *50*(4), 1029-1054.

Head, A., Mattina, T., & Smith, G. (2002). *Real exchange rates, preferences, and incomplete markets: Evidence, 1961-2001*. Unpublished manuscript, Queen's University, Canada.

Kollmann, R. (1995). Consumption, real exchange rates and the structure of international asset markets. *Journal of International Money and Finance, 14*(2), 191-211.

Lane, P., & Milesi-Ferretti, G. M. (2000). The transfer problem revisited: Net foreign assets and the real exchange rates (Working Paper No. 123). Washington, D.C.: International Monetary Fund.

Lane, P., & Milesi-Ferretti, G. M. (2001a). The external wealth of nations: Measures of foreign assets and liabilities for industrial and developing countries. *Journal of International Economics, 55*(2), 263-294.

Lane, P., & Milesi-Ferretti, G. M. (2001b). *Long-term capital movements* (Working Paper No. 8366). Boston: National Bureau of Economic Research.

Meese, R., & Rogoff, K. (1983). Empirical exchange rate models of the 1970's: Do they fit out of sample? *Journal of International Economics, 14*(1-2), 3-24.

Obstfeld, M. (1989). How integrated are world capital markets? Some new tests (Working Paper No. 2075). Boston: National Bureau of Economic Research.

Obstfeld, M., & Rogoff, K. (2000). The six major puzzles in international macroeconomics: Is there a common cause? (Working Paper No. 7777). Boston: National Bureau of Economic Research..

Rabanal, P., & Tuesta, V. (2006). Euro-Dollar real exchange rate dynamics in an estimated two-country model: What is important and what is not (Working Paper No. 177). Washington, D.C.: IMF.

Ravn, M. (2001). Consumption dynamics and real exchange rates (Discussion Paper Series No. 2940). London: CEPR.

Selaive, J., & Tuesta, V. (2003). Net foreign assets and imperfect pass-through: The consumption real exchange rate anomaly (International Finance Discussion Paper No. 764). Washington, D.C.: Board of Governors of the Federal Reserve System.

Stock, J., & Watson, M. (1996). Evidence on structural instability in macroeconomic time series relations. *Journal of Business and Economic Statistics, 14*(3), 11-30.

Stock, J., & Wright, J. (2000). Testing for weak identification. *Econometrica, 6*(3), 557-586.

Stock, J., Wright, J., & Yogo, M. (2002). A survey of weak instruments and weak identification in generalized methods of moments. *Journal of Business and Economic Statistics, 20*(4), 518-529.

Yogo, M. (2004). Estimating the elasticity of intertemporal substitution when instruments are weak. *Review of Economic Statistics, 86*(3), 797-810.

Footnotes

1 The stationarity and tractability problems associated with these models may have been the main reason.

2 Chari, Kehoe, and McGrattan (2002) referred to the discrepancy between the risk-sharing implications of theoretical models and the data as *the consumption-real exchange rate anomaly*. In Selaive and Tuesta (2003), we suggested the need for imperfect financial integration and nontradable goods to solve this anomaly. Previously, Benigno (2001) had analyzed the welfare implications of monetary policy rules under imperfect and incomplete international financial markets.

3 Lane and Milesi-Ferretti (2001b) highlighted that external wealth plays a critical role in determining the behavior of trade balance and provided some evidence that a portfolio balance exists: Real interest rate differentials are inversely related to NFA positions.

4 The consumers in both economies can trade contingent, one-period, risk-free nominal bonds.

5 This asset market structure without further modification implies an equilibrium is not well defined.

6 In log-linear form, this expression, where a caret denotes the deviation from the steady state of the log of the variable, reads as follows:

$$E_t\left(\widehat{q}_{t+1}-\widehat{q}_t\right)=E_t\left[\left(\widehat{U}_c(C^*_{t+1})-\widehat{U}_c(C_{t+1})\right)-\left(\widehat{U}_c(C^*_t)-\widehat{U}_c(C_t)\right)\right]$$

7 Chari et al. (2002) pointed out that this result stems from the fact that wealth effects in their incomplete asset market structure were very small.

8 The preferences of a household h in the country *H* are assumed to be the following: See Selaive and Tuesta (2003) for the setup used to derive these conditions and details of the well-defined steady state around which we log-linearized.

9 Some restrictions on $\phi(.)$ are necessary: $\phi(0)=1$ assumes the value 1 only if $B_{F,t}=0$, differentiable, and decreasing in the neighborhood of zero.

10 When the UIP relation holds a regression of exchange rate, returns on the interest rate differential should give an intercept of zero and a slope coefficient of unity. However, this hypothesis has been consistently rejected in the data.

11 After log-linearizing, $\delta \equiv -\varphi'(0)C$.

12 Rabanal and Tuesta (2006) performed a structural estimation with a NOEM model in which the presence of habit persistence and preference shocks augments the risk-sharing condition.

13 We used the current account and/or the GDP as the *related* series.

14 We defined the relative consumption growth rate as the logarithm of the first difference of the consumption of country *j* minus the logarithm of the first difference of the consumption of the rest of the world. We obtained the consumption of the rest of the world by aggregating the consumption of the Euro Area, Canada, Japan, and the United States and subtracting the consumption of the corresponding country *j*.

15 Backus and Smith (1993) also reported the consumption correlations against the standard deviation of the bilateral real exchange rate and found no clear role of the real exchange rate in explaining the lack of risk sharing.

16 This may be evidence that the UIP holds at longer horizons (see Chinn & Meredith, 2005).

17 In a sensitivity analysis, Ravn (2001) examined whether nonseparabilities in the utility function, aggregations over different types of goods, and habit persistence may be important in explaining risk sharing.

18 Kollmann (1995) tested the bond economy for some OECD countries and uncovered little support in the line of previous works.

19 Lane and Milesi-Ferretti (2000) argued that a model with only tradable goods might neglect the potential impact on transfers from the relative price of nontraded goods.

20 Chinn and Meredith (2005) used constant-maturity 5-year yields as a proxy for long maturities.

21 For half of the sample countries, we dealt with unit root series, which may cast doubts on some of the results.

22 $\hat{X}$ stands for log deviations around a well-defined steady state, and Δ stands for the first difference operator.

23 Empirical evidence presented by Yogo (2004) illustrates the value of the elasticity of intertemporal substitution—inverse of the risk-aversion parameter in our setup—below one.

24 Rabanal and Tuesta (2006) found values between 0.007 and 0.0129, depending on the model specification.

25 The *S-set* consists of parameter values at which one fails to reject the joint hypothesis that the parameters are the true values and that the overidentifying conditions are valid. It contains all parameters that pass the 90% χ_k^2 test, where k is the degree of freedom and, therefore, contains the topology of the objective function. As a rule of thumb, if the *S-sets* are unreasonably large, the parameters are poorly identified. See Stock and Wright (2000) for further details.

26 A proper correction of the standard errors may be appealing when the series are very persistent. However, so far, we have not found any method that could solve this problem under GMM estimation.

27 Bergin (2004) found a value of 0.0038 for the G7 countries, and Lane and Milesi-Ferretti (2001a) reported a value of 0.0254.

28 The UIP is generally tested jointly with the assumption of rational expectations in exchange markets.

29 In a seminal paper, Meese and Rogoff (1983) discovered that the predictions of a random walk dominated those of their regressions based on fundamentals for three major currencies at 6- and 12-month horizons. Please note that in this section, our intention was not to assess how economic fundamentals, in particular *NFA*, predict exchange rates.

* We are grateful to Pierpaolo Benigno, David Bowman, Mark Gertler, Joseph Gagnon, Sydney Ludvigson, Fabrizio Perri, and John H. Rogers for their suggestions and comments. We thank seminar participants at the International Finance Seminar of the Board of Governors of the Federal Reserve System 2003 for their suggestions, and Lisa Bubear for editorial assistance. We also acknowledge Jonathan Wright for facilitating some Matlab routines and for his valuable econometric advice. The opinions expressed in this paper are those of the authors and should not be attributed to the International Monetary Fund or the Banco Central de Reserva del Perú.

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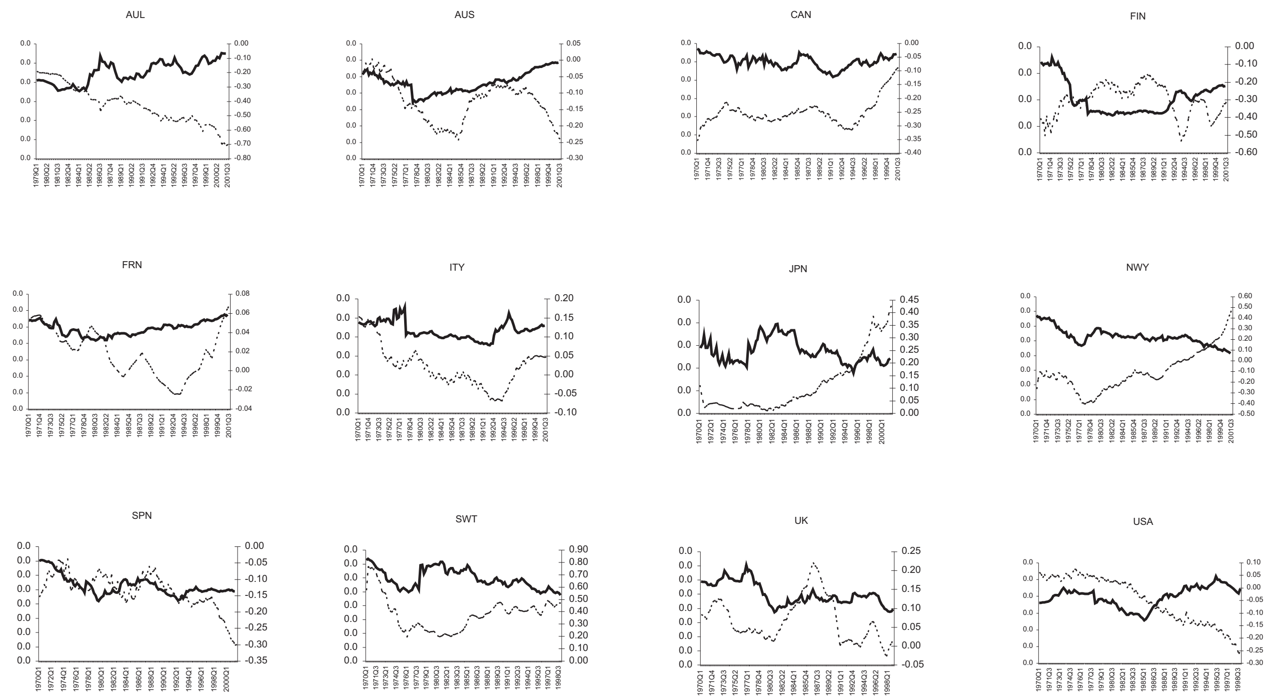
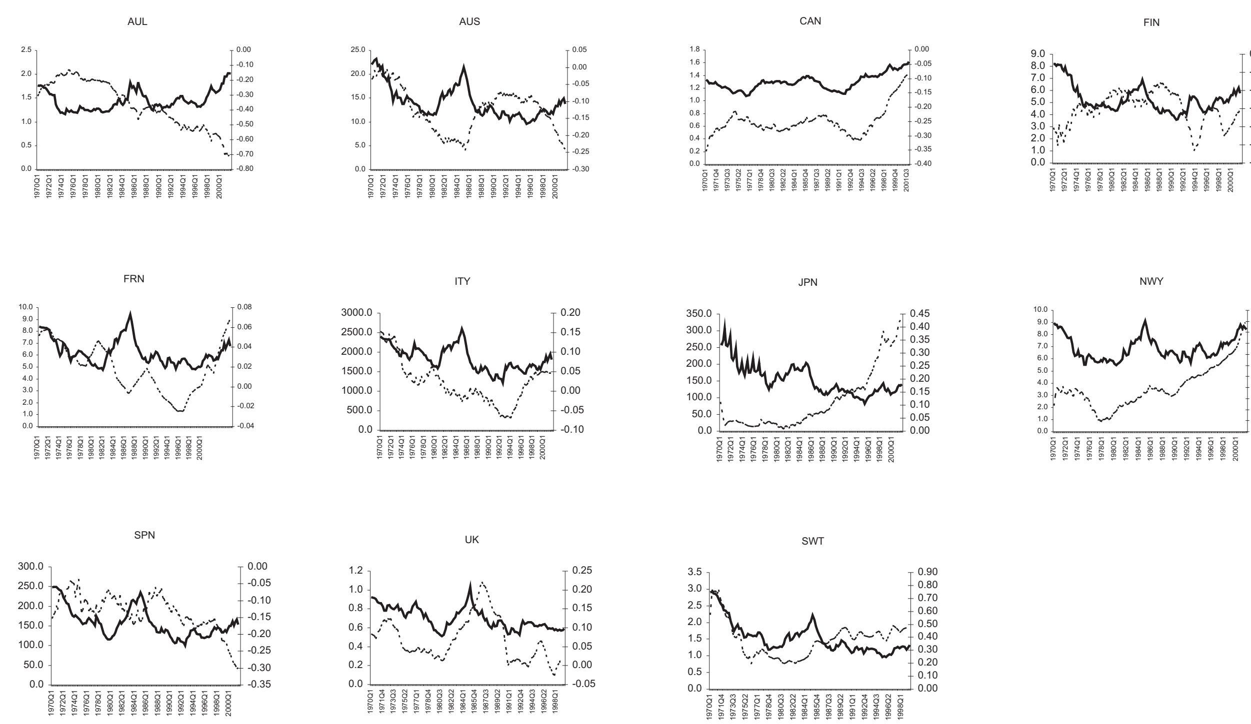
Appendix

Figure 1. Real effective Exchange Rate vs Net Foreign Assets*



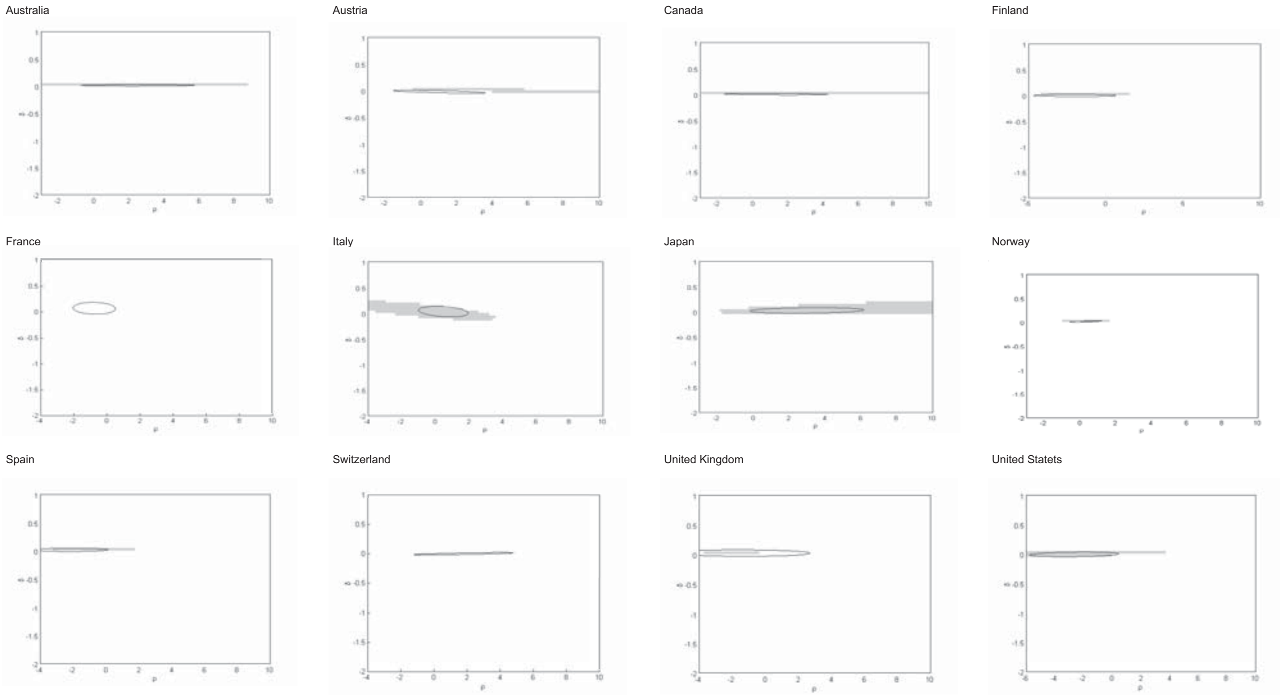
*Bold line RER, dotted line NFA.

Figure 2. Bilateral Real Exchange Rate vs Net Foreign Assets*



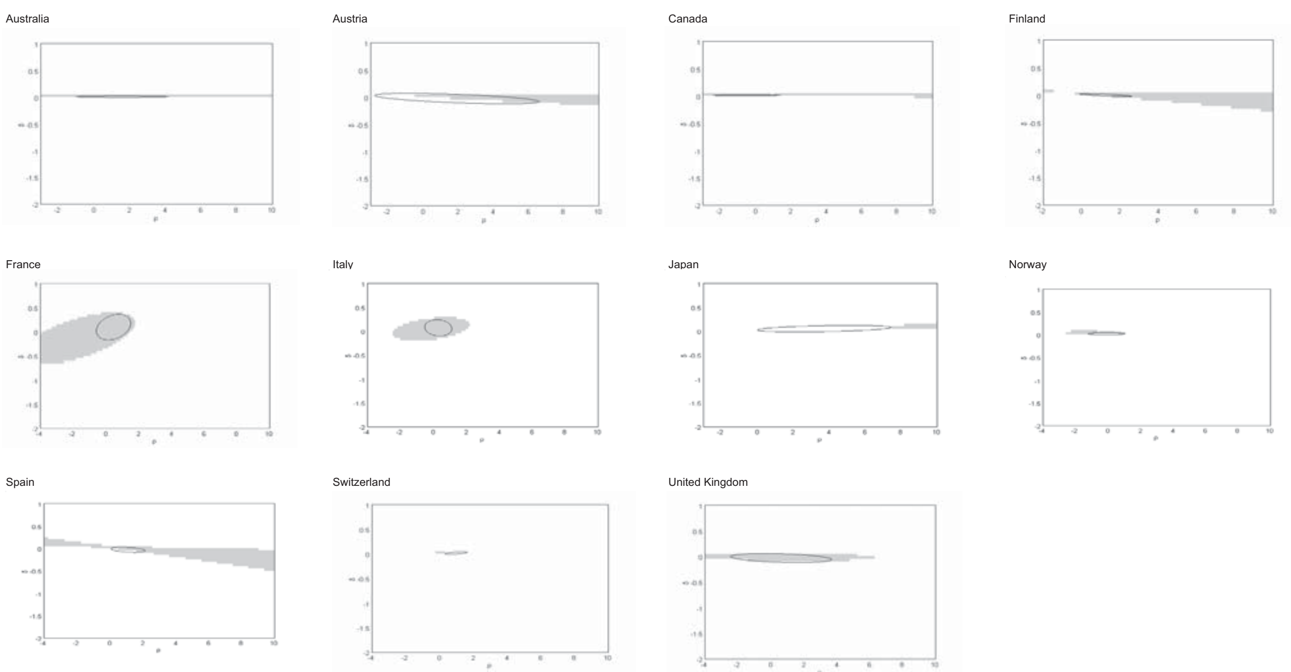
*Bold line RER, dotted line NFA.

Figure 3. S-Set for the Risk Sharing Estimation under the Rest of the World Basis



Note: Joint S-set (shaded) and 90% confidence ellipse.

Figure 4. S-Set for the Risk Sharing Estimation under Bilateral Basis



Note: Joint S-set (shaded) and 90% confidence ellipse.