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NET FOREIGN ASSETS AND IMPERFECT FINANCIAL INTEGRATION: AN EMPIRICAL APPROACH

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NET FOREIGN ASSETS AND IMPERFECT FINANCIAL INTEGRATION: AN EMPIRICAL APPROACH

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Resumen
La evidencia empírica ha rechazado de manera consistente la paridad descubierta de tasas de interés y la existencia de una alta correlación de los consumos de los países. Este trabajo investiga la importancia de mercados financieros imperfectamente integrados en estos dos temas. Bajo estos mercados, se propone una estructura donde la condición que relaciona consumos y tipo de cambio real junto a la paridad de tasas se ven afectadas por la Posición de Inversión Internacional (PII) del país. Primero, encontramos evidencia para algunos países de la OECD que la PII contribuiría a explicar la falta de correlación de los consumos. Asimismo, en términos de la paridad de tasas, la PII es capaz de capturar un premio por riesgo para un pequeño grupo de países en el corto plazo.

Abstract
Empirical evidence against both risk-sharing across countries and the uncovered interest rate parity (UIP) condition has been extensively documented. This paper investigates the empirical implications of imperfectly integrated financial markets resulting from these two issues. Under this asset market structure both the risk-sharing condition and the UIP are affected by the Net Foreign Assets Position (NFA) of the country. First, we find strong evidence for OECD countries that the NFA contributes to explaining the lack of risk-sharing across countries. Similarly, in terms of the UIP, the NFA is able to capture a time-varying risk-premium for a small group of countries over short-term horizons.

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1 Introduction

Two important issues in international macroeconomics are the apparent lack of risk sharing across countries and the UIP failure.[1] However, models with integrated financial markets and complete markets lead to 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.[2]

Regarding risk-sharing across countries, previous work has 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 (see e.g. Backus and Smith (1993), Kollmann (1995) and Obstfeld (1989)). Recently, Ravn (2001) and Head, Mattina and Smith (2002), in the line of previous findings, present evidence that an exogenous incomplete asset market structure, hereafter *bond economy*, is not supported empirically. Thus, the real exchange rate would not play a role in explaining the risk-sharing across countries. This evidence questions the empirical plausibility of recent theories of international business cycles that associate a key and significant role to the real exchange rate in breaking the link of consumption across countries.

On the other hand, it is widely accepted that the hypothesis that interest rate differentials are unbiased predictors of the nominal exchange rate performs poorly in the data.[3] Froot and Thaler (1990), in an extensive empirical testing, find striking evidence against the UIP. More recently, Chinn and Meredith (2002), hereafter CM, report that the UIP does not hold over short horizons, and present evidence suggesting that it may hold over long horizons. Bekaert, Wei and Xing (2002) find that one reason for Chinn and Meredith's claim that UIP holds better at longer horizons is simply a sample choice.

Recent theoretical studies have started to assign an explicit role to the current account and the net foreign asset position (NFA) in the transmission mechanism of shocks across countries, after being relegated to a secondary role in previous developments.[4] Selaive and Tuesta (2003) examine the role played by

[1] Obstfeld and Rogoff (2000) list the risk-sharing puzzle among the central unresolved puzzles in international macroeconomics.

[2] Scholars have incorporated an exogenous risk-premium to explain the UIP failure.

[3] The coefficient on interest rate differentials in exchange rate prediction equations turns out to be negative and significant unlike the unitary value that theory predicts.

[4] The stationarity and tractability problems associated with these models may had been the main reason to do so.

the NFA in breaking the link between real exchange rate and relative consumption. Their main theoretical contribution is that under the assumption of imperfect financial integration the NFA plays a role in explaining the apparent lack of risk-sharing across countries.[5] Cavallo and Ghironi (2002), in an overlapping generation model, try to rationalize the role of the NFA in explaining the permanent US´ real exchange rate appreciation.

From an empirical perspective, Gagnon (1996) reports a robust long-run relationship between real exchange rate and NFA in a panel of twenty $OECD$ economies. More recently, Lane and Milesi-Ferretti (2000, 2001a, 2001b) analyze the determinants of the NFA for a large set of economies, and offer a variety of theoretical reasons for thinking that some macro variables should have effects on the NFA. They also provide evidence that the NFA matters in determining long-run real interest rate differentials.[6]

The goal of this paper is to investigate the importance of the net foreign asset position in the lack of risk sharing across countries and UIP failure. We test the imperfect and incomplete asset market structure, following closely Selaive and Tuesta (2003) and Benigno (2001), in which the NFA plays a crucial role in breaking the link between the real exchange rate and the ratio of marginal utilities, and becomes a time-varying risk-premium in the UIP condition.

Two risk-free one period nominal uncontingent bonds are traded, and a cost of undertaking positions in the international financial markets allows us to characterize an imperfect and incomplete asset market structure. Under this asset market structure, and assuming deviations from "purchasing power parity" (PPP), the NFA breaks the link between the real exchange rate and relative consumptions that characterize models under complete markets. This result arises simultaneously with a direct effect of the NFA in both the UIP and the risk-sharing. 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.

[5] Chari, Kehoe and McGrattan (2002) refer 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 suggest the need of imperfect financial integration in order to solve this anomaly. In a previous contribution, Benigno (2001) analyses the welfare implications of monetary policy rules under imperfect and incomplete international financial markets.

[6] Lane and Milesi-Ferretti (2001b) highlight that external wealth plays a critical role in determining the behavior of trade balance, and also provide some evidence that a portfolio balance exists: real interest rate differentials are inversely related to the net foreign asset positions.

In terms of the risk-sharing condition, our findings suggest that growth factors of consumption and real exchange rates may behave in a manner that is consistent with a significant role for the NFA. We find relatively strong evidence in favor of a risk-sharing relationship that gives an explicit role to the NFA. For a large sample of countries, the NFA captures the smooth consumption possibilities bridging the long lasting gap between theory and data that had characterized previous works.

Regarding the UIP relationship, our findings suggest that the NFA can properly capture a time-varying risk-premium only for a small group of countries, and also allows us to obtain favorable results in terms of the unbiasedness hypothesis, i.e., the interest rate differentials are useful as predictors of short-term movements in exchange rates.

The rest of the paper is organized as follows. In Section 2 we briefly review the theoretical approach on which we based our testing. In Section 3 we present some features of the data. In Section 4 we provide an empirical discussion of both the risk sharing and UIP relationships. In Section 5, we discuss the econometric issues involved in the estimation. Section 6 provides the results, and the last section concludes.

2 A Theory of Imperfect Financial Integration

In this section, we briefly present the incomplete asset markets structure that Chari, Kehoe and McGrattan (2002), hereafter CKM, used in their work - also known as *bond economy.* Markets are perfectly integrated under CKM´s asset structure. Then, we characterize an incomplete and imperfect financial assets market structure where the NFA enters explicitly in the risk-sharing relationship and also generates deviations from the UIP.

2.1 Incomplete Markets

2.1.1 The Standard Approach: Bond Economy

It is well known that under both domestic and international complete markets, the ratio of marginal utilities of the two economies equalizes the real exchange rate[7]

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

[7] The consumers in both economies can trade contingent one-period risk-free nominal bonds.

where k_o is a function of predetermined variables, and $q_t \equiv \frac{S_t P_t^*}{P_t}$, with S as the nominal exchange rate, P^* as the Foreign price index, and P as the domestic price index. From (1), we see that the relative consumption across countries is proportional to the real exchange rate.

On the other hand, several studies on international business cycles have introduced an incomplete asset market structure in which the only asset traded internationally is a single uncontingent nominal bond.[8] Under this asset structure the risk-sharing condition reads as follows (see CKM for further details):

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

From the above expression the relation between the real exchange rate and marginal utilities holds in expected first differences.[9] As equation (2) illustrates, the *bond economy* allows us to break 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 it.[10] Furthermore, deviations from the UIP are inhibited. On the empirical grounds, evidence from Obstfeld (1989), Ravn (2001) and Head et al (2002) has cast doubts on the validity of the *bond economy* approach used by CKM.

2.1.2 Incomplete and Imperfect Financial Integration

This section follows closely Selaive and Tuesta (2003). In order to break the monotonic relationship between the real exchange rate and relative consumptions we also generate deviations from the UIP. We assume that these deviations stem from a cost of holding foreign bonds that allows us to introduce the NFA dynamics into the UIP. We may rationalize deviations from PPP either by deviations from the law of one price or by the presence of nontraded goods.

The conditions characterizing the allocations of domestic and foreign con-

[8] This asset market structure without further modification implies a non-stationary distribution of wealth across countries. Therefore, the long-run equilibrium is not well defined.

[9] In log-linear form, this expression reads as

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

where a caret denotes the deviation from the steady state of the log of the variable.

[10] CKM pointed out that this result stems from the fact that wealth effects in their incomplete asset market structure are very small.

sumption, and holding of nominal bonds are:[11]

$$
\begin{aligned}
U_c(C_t) &= (1+i_t)\,\beta \mathrm{E}_t\left\{U_c(C_{t+1})\frac{P_t}{P_{t+1}}\right\} &(3)\\
U_c(C_t^*) &= (1+i_t^*)\,\beta \mathrm{E}_t\left\{U_c(C_{t+1}^*)\frac{P_t^*}{P_{t+1}^*}\right\} &(4)\\
U_c(C_t) &= (1+i_t^*)\,\phi\left(\frac{B_{F,t}S_t}{P_t}\right)\beta \mathrm{E}_t\left\{U_c(C_{t+1})\frac{P_tS_{t+1}}{P_{t+1}S_t}\right\} &(5)
\end{aligned}
$$

where β is the intertemporal discount factor, and $\phi(.)$ depends on the real holdings of the foreign assets in the entire economy, and therefore is taken as given by the domestic household.[12]

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 are able to derive the new uncovered interest parity and the risk-sharing equilibrium condition. Both are affected by the net foreign asset position of the domestic economy.

The uncovered interest rate parity is derived by taking the difference between the log-linear approximation of equations (3) and (5), and is given by the following expression:

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

Notice that the above equation incorporates a cost of borrowing in foreign currency and may be consistent with the empirical failure of the UIP.[13] In our case, there is a time varying risk-premium that depends on both the NFA of the domestic economy, b_t, and a cost of bond holdings, δ, that measures the elasticity of the interest rate differential to changes in the NFA position.[14] The higher this 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

[11] The preferences of a household h in the country H are assumed to be $U_t^h = E_t\left\{\sum_{s=t}^{\infty}\beta^{s-t}\left[U(C_{t+s}^h) + L\left(\frac{M_{t+s}^h}{P_{t+s}}\right) - V\left(N_{T,t+s}^h, N_{NT,t+s}^h\right)\right]\right\}$. See Selaive and Tuesta (2003) for the set up used to derive these conditions, and details of the well defined steady state around which we log-linearize.

[12] 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.

[13] 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.

[14] After log-linearizing, $\delta \equiv -\phi'(0)\,\overline{C}$.

depending on the Home country being a borrower or a lender in the international assets market. Observe that this equation implies a negative relation between the interest rate differential and the NFA of the economy.

The *risk-sharing* condition under the imperfect financial integration we impose here is obtained by combining the UIP and the corresponding Euler equations for each country, and reads as:

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

Equation (7) illustrates the mechanism through which the NFA position affects the risk-sharing. The characterization of this incomplete asset market structure maintains the gap between relative consumptions that emerges in the *bond economy* specified in equation (2), but now, in addition, the dynamic of the NFA plays an explicit role. As long as there is either asset accumulation or decumulation, the real exchange rate will be affected by the NFA, and the link between the real exchange rate and relative consumptions will be broken down. Ceteris paribus, there is a negative relation between the real exchange rate and the NFA, i.e., an asset accumulation implies a real exchange rate appreciation. The larger the asset accumulation the greater will be 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 $\delta \to 0$ or $b_t = 0$ at every period, the risk-sharing relationship boils down to the *bond economy.*

3 Features of the Data

3.1 Data

All the data collected in the paper corresponds to quarterly series with the exception of the net foreign asset position that is available only at annual frecuency from 1973 to 1998. The series of consumption correspond quarterly series of private non-durables final consumption at constant prices, and were obtained from the OECD's Quarterly National Accounts (QNA) and IMF's International Financial Statistics (IFS). The series were deflated by the corresponding implicit price deflator for final consumption, and then multiplied by the nominal exchange rate to express them in terms of US constant dollars.

The series of NFA positions were obtained from Lane and Milesi-Ferretti (2001a)'s database. The NFA were interpolated to get quarterly series by Chow and Lin (1971)'s methodology.[15] The variable was scaled by the GDP in current dollars of the corresponding year. We complete the data for the period 1999 to 2001 using the quarterly cumulative current account.

Bilateral real exchange rates are defined as the nominal exchange rate times the ratio of foreign to domestic prices. Nominal exchange rates were obtained from the IFS, and prices are defined as the implicit deflactors for the consumption variables.

The real effective exchange rate is obtained from the IFS for the period 1975.1-2001.4. In order to complete back the sample until 1973 we define the real effective exchange rate as the nominal effective exchange rate times the ratio of the aggregate OECD prices to domestic prices. The nominal effective exchange rates are taken from the IFS.

The interest rate corresponds to 3-months and 12-months euro-currency yields expressed in annual terms, and were obtained from the Bank of International Settlements´ database.

3.2 Description of the Data

Under the standard assumption of separability in the utility function and allowing for PPP, a complete asset market assumption will imply a perfect cross-correlation of consumptions across countries. In Table 1 we perform the cross-country correlations of consumption growth rates. For most of the countries this correlation is very low, and gets higher when it is calculated with respect to the an aggregate of OECD consumption, consistent with Ravn's findings.

As we pointed out in the previous section, the real exchange rate introduces a wedge between the relative consumptions across countries, and therefore, 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. In Table 2 we calculate cross-correlations between bilateral (and effective) real exchange rates and consumption growths rates for twelve OECD economies for the period 1970.1- 2001.4.[16] These cross-correlations are quite low and negative in most of

[15] We use the Current Account and/or the GDP as the *related* series.

[16] We define 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 cases, so it seems that the *bond economy* may not be supported by the data. Econometric estimations by Ravn (2001) and Head et al (2002) also confirm this finding.[17]

The theory we test in this paper considers the NFA as a key determinant of the lack of risk-sharing across countries (see equation 7). Figures 1 and 2 plot, for the period 1973-2001, the NFA position vis a vis the real effective and bilateral exchange rate, respectively. The bold line represents the real exchange rate and the dotted line stands for the NFA. Most of the real exchange rate series exhibit large swings around a slowly drifting mean. The NFA drifts upward for Japan, Norway and Switzerland, and downwards for Australia and United States, with little trend in the remaining countries. The theory we test in this work would predict a positive correlation between the expected growth rate of the real exchange rate and NFA.

Table 2 reports the correlations between the growth rate of the real exchange rate and the NFA for the whole set of countries. The results are mixed with positive and negative correlations, and most of them quite low. This is preliminary evidence in favor of a theory in which the NFA may play a role for some economies.

We perform a similar exercise for the UIP condition. In Table 3, we present 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 when we use a short-term interest rate differential, although they increase when we use the 12-months interest rate differential.[18] On the other hand, the NFA position is positively correlated with the expected change in the exchange rate for 6 out of 14 countries.

The above evidence is suggestive indicating 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 a robust empirical testing.

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the rest of the world. Consumption of the Rest of the World (RoW) is obtained by aggregating the consumption of the Euro Area, Canada, Japan and US, and substracting the consumption of the corresponding country j .

[17] Backus and Smith (1993) also report the consumption correlations against the standard deviation of the bilateral real exchange rate, and find no clear role of the real exchange rate in explaining the lack of risk sharing.

[18] This may be evidence that the UIP holds at longer horizons (see Chinn and Meredith (2002))

4 Empirical Discussion

Obstfeld (1989) first derived and tested a risk sharing hypothesis in a set up where PPP did not hold. Ravn, following a similar approach, also assumed that countries can borrow and lend freely at the same nominal interest rate. Both authors do not find evidence in favor of the *bond economy* setup.[19] Head et al (2002) extended these previous works 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 do not support any of these extensions.

Our econometric approach will follow Kollmann (1995) who studied the relation between consumption and the real exchange rate using the Generalized Method of Moments estimation procedure (GMM).[20] Although, our approach differs from Kollmann's in several dimensions. First, we investigate a broader data set. Second, we test a different risk sharing condition in which the NFA enters explicitly. Third, we also test the UIP condition that arises from the asset market structure we suggest in section 2.

There are some reasons to believe that the current account indeed plays an important role in the international transmission mechanism of shocks. Gagnon (1996) present evidence of a significant and robust long-run relationship between the real exchange rate and the NFA. Recently, Lane and Milesi-Ferretti (2000) argue that the NFA has a strong impact on the relative price of non-traded goods, and therefore, on the real exchange rate dynamics.[21]

With respect to the UIP, the hypothesis that interest rate differentials are unbiased predictors of future exchange rate movements has been extensively rejected in empirical studies. 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-horizons UIP

[19] In his sensitivity analysis, Ravn examines whether non-separabilities in the utility function, aggregations over different types of goods, and habit persistence may be important in explaining the risk sharing.

[20] Kollmann tests the *bond economy* for some OECD countries, and finds little support for it in the line of posterior works.

[21] These authors argue that a model with only tradable goods may neglect the potential impact on transfers from the relative price of non-traded goods.

holds. Recently, CM tested the UIP hypothesis on longer-maturity bonds for US, Germany, Japan and Canada, and find evidence that the longer the maturity the better the interest rate differential does explaining the future exchange rate variations.[22] They interpret this as meaning that any risk premium is very stable over long horizons. Although, Bekaert et al (2002) find that one reason for Chinn and Meredith's claim that UIP holds better at longer horizons is simply a sample choice. Finally, Lane and Milesi-Ferretti (2001b) find a strong long-run link between the real interest rate differential and the NFA for a large sample of countries.

5 Econometric Issues

5.1 Persistence of the 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 and Watson (1996)).

On the other hand, we deal with a sample period that goes back until 1973.1 which limits the number of observations to no more than 100. In this context, to rely in tests of the null hypothesis of a stationary process -as the KPSS and LMC test- that are based on conventional asymptotic critical values may mislead to reject the null hypothesis. Conversely, Caner and Kilian (2001) have shown that tests that rely on the null hypothesis of unit root may overcome this problem when one corrects the critical values for finite-sample or bootstrap critical values.

Thus, we use the efficient *DF-GLS* test (Elliott et al (1996)) of the unit root null hypothesis using finite-sample critical values. We follow closely Caner and Kilian (2001) to create finite-sample critical values that we compare with the statistic generated by the test. The results are reported in Table 4. After applying the test to the NFA for the set of countries in our sample, we reject the unit root null at the 10 percent for 7 out of 14 countries. The previous result suggests that the NFA is not only theoretically, but also empirically mean-reverting for some countries, and therefore, it is plausible that countries

[22] These authors use constant-maturity 5-year yields as a proxy for long maturities.

that cannot reject the null of unit root are strongly influenced by a small sample size problem.[23]

5.2 Estimation Procedure

We will use the GMM procedure developed by Hansen (1982). Under this estimation procedure, we minimize a criterion function that is derived by imposing at least as many moment conditions as parameters to be estimated.

It is well known that if the instruments are poorly correlated 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. This problem cannot be avoided by enlarging the sample or increasing the number of instruments as Stock, Wright and Yogo (2002) point it out. Despite the evolving nature of this literature, there are some useful methods to address concerns about weak identification. Recently, Stock and Wright (2002) developed a test to check for the presence of weak instruments. We will follow these later authors, and implement a fully robust test for weak identification (*S-set test*). Under this test, we use the robust continuous-updating GMM estimator, which minimizes the following objective function:

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

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

We will juxtapose the conventional 90% confidence ellipse with the 90% *S-Set*. This *S-Set* contains all parameters that pass 90% χ^2_k test, and is constructed according to Theorem 2 in Stock and Wright. Loosely speaking, we would not have weak instruments if the *S-Set* is contained in the 90% confidence ellipse.

[23] The fact that for half of the sample of countries we deal with unit root series may cast doubts about some of our results.

6 Results

6.1 Testing a Risk-Sharing condition under Imperfect Financial Integration

Our incomplete and imperfect asset market structure delivers the following orthogonality condition to be estimated:[24]

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

where 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 examine the risk sharing condition, equation (8), using quarterly data for a set of 12 countries. We perform three estimations. First, we consider the real effective exchange rate and relative consumption growth rates with respect to the RoW. Second, we examine the same risk sharing in a country-pair basis with respect to US. Finally, we perform a balanced panel for both cases.

The results of the estimation country-RoW are reported in Table 5. By way of contrast, in the second and third columns we present the estimation that corresponds to the *bond economy*. The estimated coefficient of risk aversion, ρ, is negative for seven of the twelve countries and is positive and significant only in two cases (Japan and Italy). The associated *p-values* of the J statistics are above 0.1 in all cases, so we do not reject the null of overidentifying restrictions. The previous results suggest, at least, weak evidence in favor of the *bond economy*.

Next, we test the risk sharing relationship proposed in the paper, equation (8). The results are in the last columns of Table 5. The first striking result is that the estimate of the risk-aversion parameter turns out to be positive and significant for seven out of twelve countries, which may suggest that the instruments associated to the NFA positions are helping to identify the risk-aversion parameter, and to capture some aspects of smooth consumption possibilities.[25] The second result to highlight is the positive and significant value of the cost of bond holding parameter, δ, for five countries in the sample. The associated

[24] $\widehat{X}$ stands for log-deviations around a well defined steady state, and Δ stands for the first difference operator. The results do not change significantly after hp-filtering the consumption and real exchange rate series.

[25] Recent empirical evidence presented by Yogo (2002) locates the value of the elasticity of intertemporal substitution -inverse of the risk aversion parameter in our set up- below one.

p-values of the J statistics are all above 0.10. Finally, it is worthwhile to notice that for three countries (Australia, Japan and Norway) both parameters, ρ and δ are positive and significant.

We examine the possibility that the previous results may be driven by weak identification problems. To do so we construct the conventional 90% confidence ellipse with the 90% S-set described before. The results are summarized in the last column of Table 5, and the *S-set test* are plotted in Figure 3.[26] Under the reasonable assumption that the risk-aversion parameter is not "too large" as previous empirical evidence has suggested (see e.g. Yogo, 2002), our estimations may not be driven by weak identification.

The panel GMM estimation is reported in the last row of Table 5. Both the coefficient of risk aversion and the cost of bond holding parameter are significant and have the correct sign. Thus, we give support for a theory of imperfectly integrated financial markets.

Finally, we turn to the estimation of the country-by-country basis. Results are reported in Table 6. Again in the second and third columns we report 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. When we include the NFA in the equilibrium condition we improve considerably the estimations. The estimate of the risk-aversion parameter turns out to be positive and significant for 8 out of 11 countries, and the estimate of the cost of bond holdings is also positive and significant for 7 countries.[27] The associated *p-values* of the J statistics are above 0.10 in all cases. In most of the cases, 8 out of 11, our results are not driven by weak identification problems as it is shown in figure 4. Again, our panel results support the theory since the estimates of parameters ρ and δ are positive and significant. The previous results may suggest that a theory of imperfect financial integration may work better in a country-by-country than in country-RoW basis.

[26] The *S-set* consists of parameter values at which one fails to reject the join hypothesis that the parameters are the true values and that the overidentifying conditions are valid. It contains all parameters that pass the 90% χ^2_k 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, then the parameters are poorly identified. See Stock and Wright (2000) for more details.

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

Overall, the tested risk-sharing condition works 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 our countries are in the line of what theory would predict. In a nutshell, it appears that growth factors of consumption and real exchange rates behave in a manner which may be consistent with the assumptions implicit in our incomplete and imperfect market structure.

6.2 Testing the Uncovered Interest Rate Parity

In this section we want to examine the role of the NFA in explaining the UIP condition in the short run rather than testing the UIP at different horizons.[28] The market structure outlined in section 2, in a regression context, delivers the following orthogonality condition:

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

where 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.

We estimate the UIP, equation (9), by GMM for two different horizons with 3- and 12- maturity bonds (as in CM, 2002). At this stage, we do not claim that the NFA may help to predict exchange rate movements, but we want to assess whether there is a significant role for the NFA in explaining exchange rate movements as our theory would predict.[29] In particular, we want to assess the significance and sign of the parameter δ in equation (9).

Using constant-maturity 3-months yields for 14 countries, we implement regressions of the form of equation (9) over the 1980.1-2001.4. As a way of contrast, we also display the estimates of the UIP condition analyzed in previous studies. We use as a benchmark the results obtained by CM. They estimate the UIP for short-term horizons -3-months and 12-months- for the Deutschemark, Japanese yen, UK pound, French franc, Italian lira and Canadian dollar. In our exercises we have used exactly the same data set as CM. The exchange rates

[28] To operationalize the convet, UIP is generally tested jointly with the assumption of rational expectations in exchange markets.

[29] In a seminal paper, Meese and Rogoff (1983) find that the predictions of a random walk dominates those of their regressions based on *fundamentals* for three mayor currencies at 6- and 12-months horizons. It is worth stressing that in this section our intention is not to assess how economic fundamentals, in particular NFA, *predict* exchange rates.

of each country were expressed in terms of the US dollars, and the 3- and 12-months movements in the exchange rate were regressed against differential in euro currency yields of the corresponding maturity. Since 12-horizon data at a quarterly frecuency may led to MA(3) in the residuals, we use the Newey-West correction to get robust standard errors.

The results reported in Table 7 present the estimations for 3-months 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 the line with the results reported by CM.[30] The estimate slope coefficient, β, has negative sign for 12 out of 14 countries. Only France and Italy present coefficients that are not statistically different than 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 remarkably result is 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 10 percent level. The previous finding stands in contrast with CM that report the failure of the UIP for most of the currencies analyzed at 3-months horizons. On the other hand, 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, we observe that the slope coefficient moves on the right direction, but the risk-premium parameter has the wrong sign and/or is positive but not significant.

We perform the same estimation for longer maturity bonds, and the results at 12-months maturity are shown in Table 8. Again, we find some support for our theory. The slope parameter is significant and has the correct sign for seven countries. These results improve slightly with respect to the 3-months horizon. The risk-premium parameter is also positive for 7 countries, and is significant for five of them.

In a nutshell, it seems that NFA may be useful predictors of short-term movements in exchange rate for some countries, and they are likely to explain the observed variance in exchange rates for the period analyzed. Even though, our findings suggest that the NFA may not be an appropriate measure of time-

[30] The results confirm the failure of the UIP similar to other studies by Froot and Thaler (1990). If UIP holds, the slope coefficient should not be statistically different that one. Regarding the constant term, non-zero values may be explained by Jensen's inequality, and are not shown to save space.

varying risk premium for an important subsample of countries.

6.3 Joint Test of the Risk-Sharing and the UIP

We also perform a tighter test of the imperfect and incomplete asset market structure presented in section 2. We implement a joint GMM estimation of the risk sharing and UIP equilibrium conditions. Under the joint estimation, we impose the risk premium parameter, δ, to be the same in both equations. One of the limitations of this approach is that the sample size is limited to the sample period used in the UIP estimation while we increase the number of moment conditions and parameters to be estimated. To do a balance estimation we restrict to 10 the number of countries in the sample. Results are reported in Table 9.

The estimated risk aversion parameter, ρ, is positive and significant for seven countries. On the other hand, the slope coefficient of the UIP condition has the right sign for 4 countries. Finally, parameter δ that intends to capture the time varying risk premium generated by the NFA is significant at 10 percent for 5 economies.

The results of these estimations seem to point out in two directions. Firstly, there is a significant role for the NFA position in explaining the lack of risk sharing across countries. Secondly, there seems to be a significant link between the UIP puzzle and the apparent lack of risk sharing across countries.

7 Concluding Remarks

This paper has looked to 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. The empirical failure of most of the theoretical models under the assumption of perfect integration, even when it is allowed for both exogenously and endogenously incomplete markets, habit persistence and different forms of utility functions, has been extensively documented. Recent evidence on the importance of the NFA in explaining the transmission of shocks across countries has suggested us to consider the implications of imperfect financial integration in international macroeconomics. The results of the paper contrast with those of previous studies based on the assumption of complete markets. Firstly, we find evidence that growth factors of consumption and real exchange rates may behave in a manner which is consistent with a significant role for the NFA for a large sample of OECD countries. In this sense, the NFA is a key element in explaining the apparent lack of risk sharing across countries. Secondly, for a small group of countries, the NFA captures a time-varying risk-premium and yields a positive slope coefficient for the interest rate differential at short-term horizons which stands in contrast with Chinn and Meredith (2002)'s findings. In this sense, the interest rate differential could be a useful predictor of short-term movements in the nominal exchange rate when it is accompanied by the NFA position.

In a nutshell, it seems reasonable to consider a theory where the NFA position affects both the risk-sharing across countries and the UIP condition. Our findings would suggest that since the NFA helps to explain nominal exchange rate movements, an important avenue to investigate is the predictability power of the NFA following Meese and Rogoff (1983)´s seminal contribution. In this line, to overcome the high persistence of the NFA in the empirical testing by a suitable transformation may be a good alternative.

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Table 1. Cross-country consumption correlations: 1973.1-2001.4

Correlation $((C^{i}_{t+1} - C^{i}_{t}, C^{j}_{t+1} - C^{j}_{t}))$

	Austria	Canada	Finland	France	Italy	Japan	Norway	Spain	Switzerland	U.Kingdom	US	OECD
Australia	0,11	0,14	0,37	0,05	0,15	0,04	-0,08	0,33	0,32	0,14	-0,03	0,15
Austria		0,02	0,00	0,00	0,20	0,05	0,04	0,33	0,21	0,04	0,02	0,17
Canada			0,06	0,11	0,28	0,06	-0,11	0,29	0,11	0,26	0,36	0,31
Finland				-0,24	0,12	0,08	0,17	0,33	0,34	0,54	0,14	0,34
France					0,00	0,23	-0,68	0,01	-0,17	0,07	0,13	0,08
Italy						0,03	0,06	0,51	0,36	0,15	-0,10	0,32
Japan							-0,17	0,10	0,05	0,17	0,17	0,22
Norway								0,04	-0,01	-0,05	0,02	-0,09
Spain									0,30	0,26	0,11	0,44
Switzerland										0,09	-0,07	0,22
U.Kingdom											0,36	0,47
US												0,28

Notes: - Series of consumption correspond to private non-durable final consumption obtained from OECD´s Quarterly National Accounts and IMF´s International Financial Statistics.

Table 2. Unconditional Correlations 1973.1-2001.4: *Risk Sharing*

Country	Correlation of the Change in Real Effective exchange rate with[a]:		Correlation of the Change in Real Bilateral exchange rate with[b]:	
	$\Delta C^{R\text{-}ROW}_{t+1,i}$	$b_{t,i}$	$\Delta C^{R\text{-}US}_{t+1,i}$	$b_{t,i}$
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
U.Kingdom	0.03	0.12	-0.19	-0.15
US	-0.10	0.01	-	-

Notes: - $\Delta C^{R\text{-}ROW}_{i}$ is the relative consumption growth rate of country *i* with respect to the Rest of the World

- $\Delta C^{R\text{-}US}_{i}$ is the relative consumption growth rate of country *i* with respect to United States (US).

- b_i corresponds to the ratio *NFA/GDP* of country *i* .

[a] We use the multilateral effective real exchange rate of country *i* .

[b] We use the bilateral real exchange rate of country *i* with respect to US.

Table 3. Unconditional Correlations 1973.1-2001.4: Uncovered Interest Parity

Country	Correlation of the Change in Nominal Exchange Rate with[a]:		
	$\Delta i_{t,i}^{R\text{-}3m}$	$\Delta i_{t,i}^{R\text{-}12m}$	$b_{t,i}$
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
U.Kingdom	-0.19	0.05	-0.15

Notes: - $\Delta i_{t,i}^{R\text{-}3m}$ and $\Delta i_{t,i}^{R\text{-}12m}$ are the 3- and 12-months interest rate differentials in euro currency yields of country *i*.

- b_i corresponds to *NFA/GDP* of country *i*.

[a] Corresponds to the change in the bilateral nominal exchange rate of country *i* with respect to US. The change in the nominal exchange rate corresponds to the same maturity of the interest rate differential.

Table 4. *DF-GLS* Test for the *Net Foreign Asset Position*

Country	*DF-GLS*	Reject I(1) null (5 or 10%) [a]	Sample Period
Australia	-1.732	*yes*	1973:1 – 2001:4
Austria	-4.562	*yes*	1973:1 – 2001:4
Canada	-1.693	*yes*	1973:4 – 1997:3
Finland	-1.879	*yes*	1973:1 – 2001:4
France	-1.715	*yes*	1973:1 – 2000:4
Germany	-0.516	*no*	1973:1 – 2001:4
Italy	-0.437	*no*	1973:1 – 2001:4
Japan	0.002	*no*	1973:1 – 2001:4
Netherlands	-0.433	*no*	1973:1 – 2001:4
Norway	0.354	*no*	1973:1 – 2001:4
Spain	-1.881	*yes*	1973:1 – 2000:4
Sweden	-0.165	*no*	1980:4 – 2001.4
Switzerland	-0.948	*no*	1973:1 – 2001:4
U.Kingdom	-1.801	*yes*	1973:1 – 2001:4

Notes: - We allowed for 8 lags to construct the statistic and the finite sample critical values.
[a] Finite-sample critical values -2.09[-1.69] at 5[10]%.

Table 5. Risk-Sharing with the Rest of the World

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

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

	(1)		(2)			
Country	ρ	**J-stat**	ρ	δ	**J-stat**	**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 [a]	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 and were 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.

*, (**), [***] Significance at 10%, (5%), [1%].

[a] All countries but USA and Australia.

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 and were modified by the 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.

*, (**), [***] Significance at 10%, (5%), [1%].

[a] All countries but USA and Australia.

Table 7. Uncovered Interest Parity: Maturity 3 months

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

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

Country	(5) Estimate β	(6) *Maturity*: 3 months Estimate β	(6) *Maturity*: 3 months 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 are reported in parenthesis and were modified by the Newey-West correction.
- All *p-values* of *J- statistics* are above 0.1.
- Bilateral Nominal Exchange Rate in terms of US dollars. Interest Rate differential in Eurocurrency yields.
*(**)[***] Different from null of unity at 10%(5%)[1%].
†(††)[†††] Different from null of zero at 10%(5%)[1%].

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	(7) Estimate β	(8) *Maturity*: 12 months Estimate β	(8) *Maturity*: 12 months 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 are reported in parenthesis and were modified by the Newey-West correction.
- All *p-values* of *J- statistics* are above 0.1.
- Bilateral Nominal Exchange Rate in terms of US dollars. Interest Rate differential in Eurocurrency yields.
*(**)[***] Different from null of unity at 10%(5%)[1%].
†(††)[†††] Different from null of zero at 10%(5%)[1%].

Table 9. Joint Estimation of the UIP and Risk-Sharing Condition.

$$\rho[C_{t+1} - C_t - (C^*_{t+1} - C^*_t)] - (q_{t+1} - q_t) + \delta b_t = 0$$

$$\alpha + \beta\,(i_{t,k} - i^*_{t,k}) + \delta b_{t,k} - \Delta S_{t,t+k} = 0$$

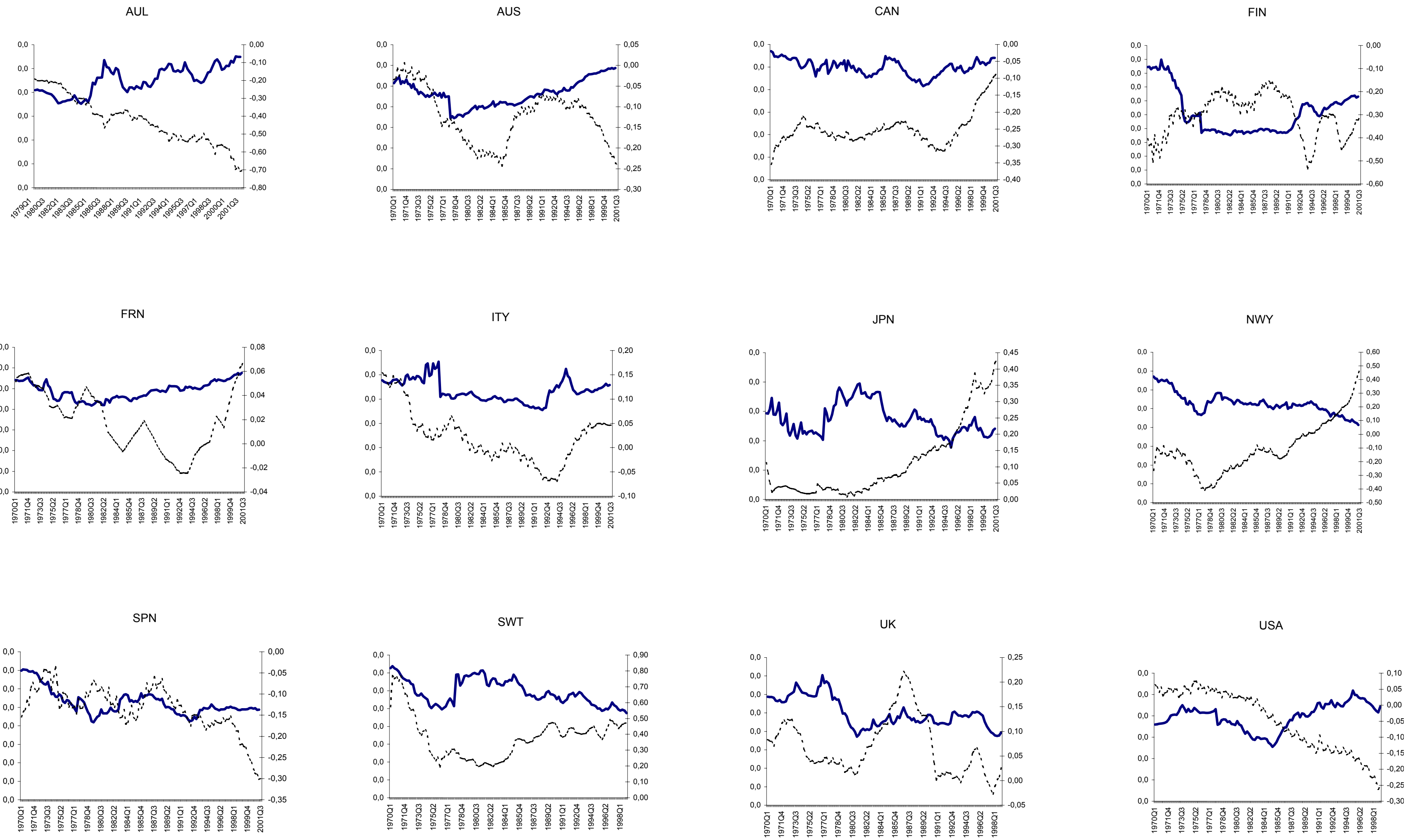
	Joint Estimation		
Country	*Estimate* ρ	*Estimate* β	*Estimate* δ
Australia	-0.254	0.445	-0.013†††
	(1.328)	*(0.401)*	*(0.004)*
Austria	3.979†	-0.742**	-0.074†††
	(2.323)	*(0.273)*	*(0.031)*
Canada	1.164†††	-0.758***	-0.007†
	(0.152)	*(0.214)*	*(0.004)*
France	-0.738†	-1.019***	0.140††
	(0.395)	*(0.670)*	*(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 are reported in parenthesis and were modified by the Newey-West correction.
- All *p-values* of *J- statistics* are above 0.1.
- Bilateral Nominal Exchange Rate in terms of US dollars. 3-months Interest Rate differential in Eurocurrency yields.

*(**)[***] Different from null of unity at 10%(5%)[1%].

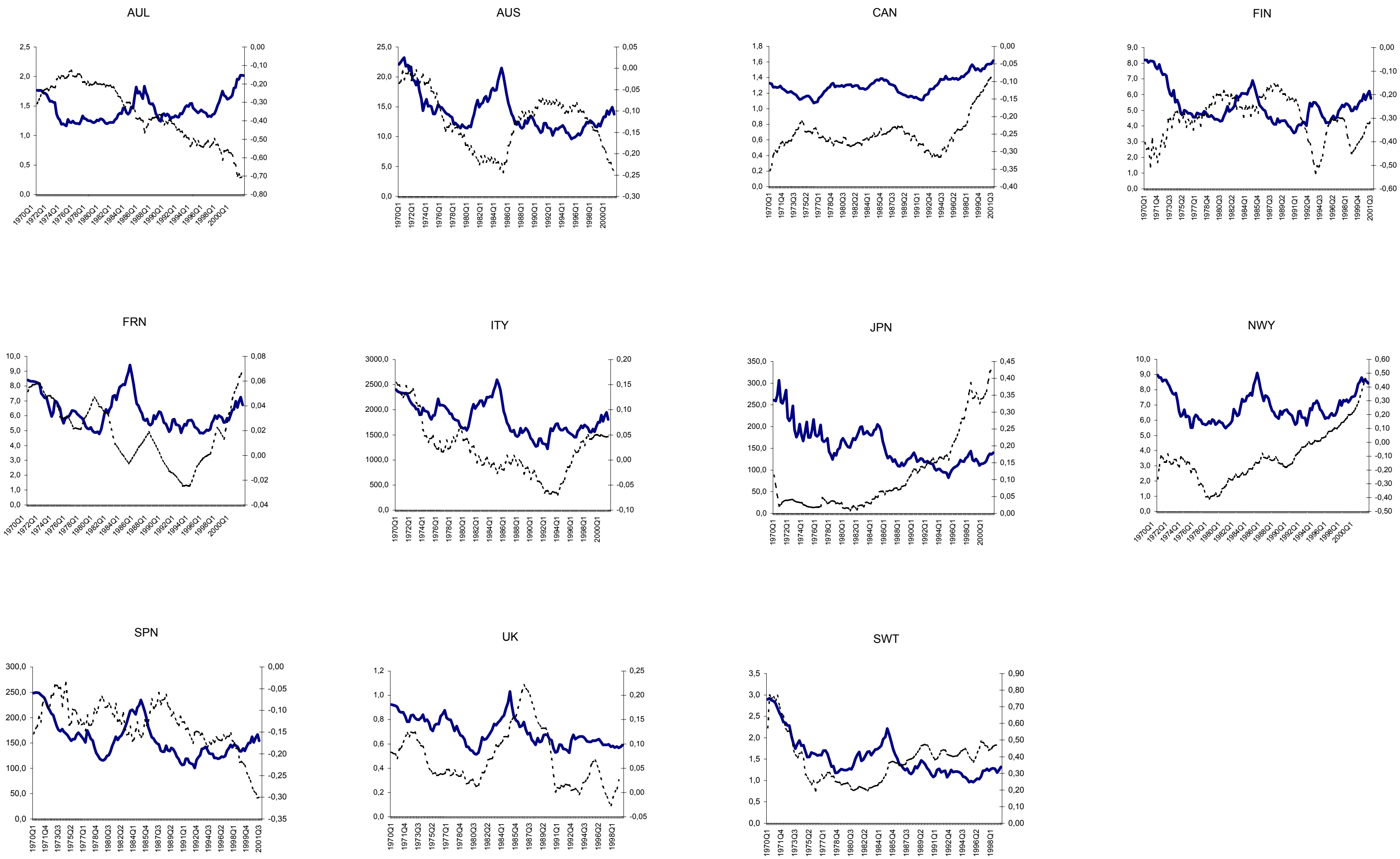
†(††)[†††] Different from null of zero at 10%(5%)[1%].

Figure1: Real Effective Exchange Rate vs Net Foreign Assets*



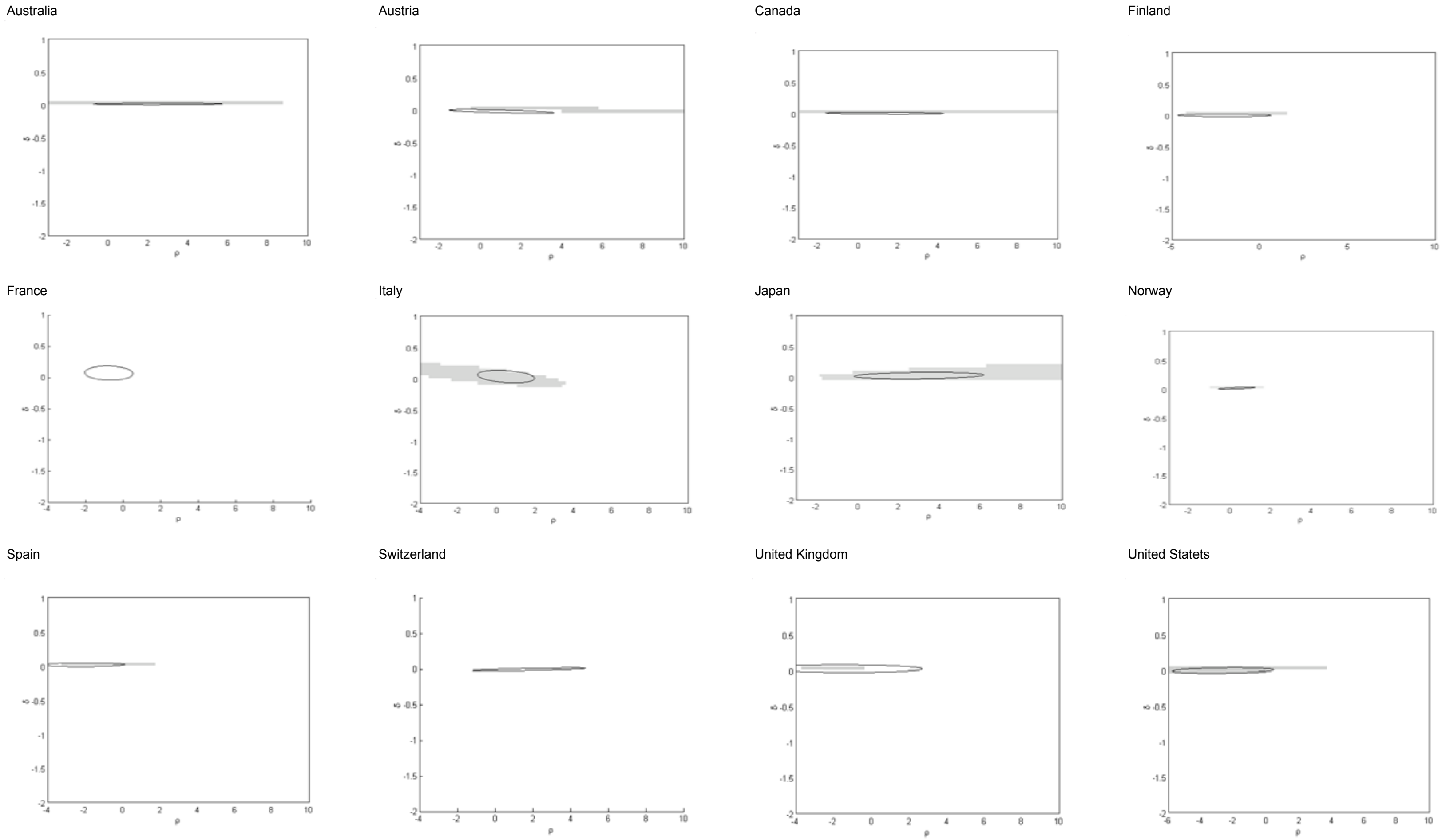
* Bold line RER, dotted line NFA.

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



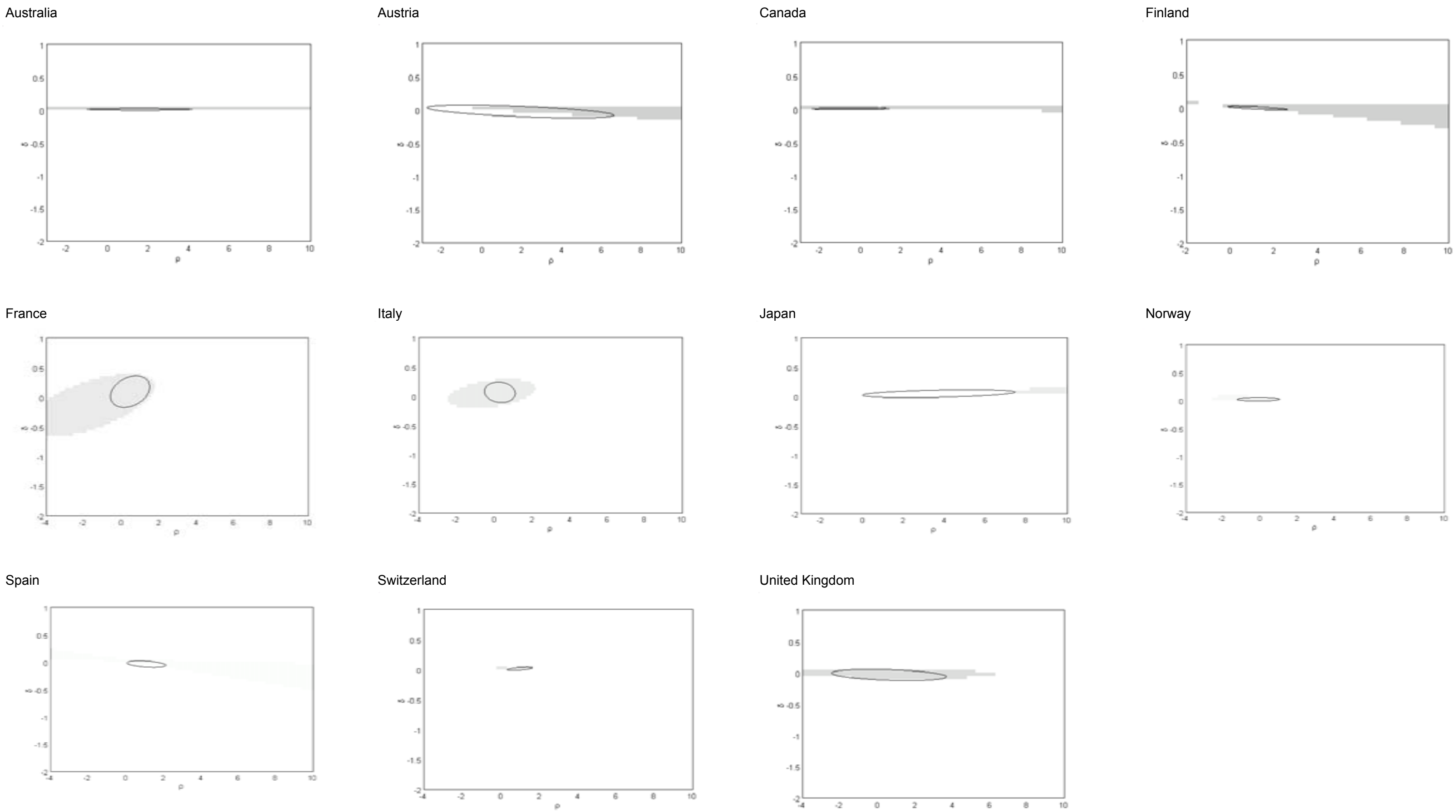
* Bold line RER, dotted line NFA.

Figure 3: S-Set for Risk Sharing with the Rest of the World Estimation



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

Figure:4 S-set for the Bilateral Risk-Sharing Estimation



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

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