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Giancarlo Corsetti, Luca Dedola and
Francesca Viani

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Giancarlo Corsetti, Cambridge University, EUI and CEPR
Luca Dedola, European Central Bank and CEPR
Francesca Viani, Bank of Spain

ABSTRACT

The International Risk-Sharing Puzzle is at Business Cycle and Lower Frequency*

We decompose the Backus-Smith [1993] statistic—a low or negative correlation between relative consumption and the real exchange rate at odds with a high degree of international risk sharing—in its dynamic components at different frequencies. Using multivariate spectral analysis techniques we show that, in most OECD countries, the dynamic correlation tends to be more negative, and significantly so, at business cycle or lower frequencies—the appropriate frequencies for assessing the performance of international business cycle models. Theoretically, we show that the dynamic correlation predicted by standard open-economy models is the sum of two terms: a term constant across frequencies, which can be negative as a function of uninsurable risk; a term variable across frequencies, which in bond economies is necessarily positive, reflecting the insurance intertemporal trade provides against forecastable contingencies. We show that the main mechanisms proposed in the literature to account for the puzzle are consistent with the evidence.

Giancarlo Corsetti
Department of Economics
Cambridge University
Sidgwick Avenue
Sidgwick Avenue
Cambridge CB3 9DD

Luca Dedola
Monetary Policy Research Division
European Central Bank
Postfach 16 03 19
D - 60066 Frankfurt am Main
GERMANY

Email: gc422@cam.ac.uk

Email: luca.dedola@ecb.int

Francesca Viani
International Economics Division
Bank of Spain
c/ Alcalá 48
28014 Madrid
SPAIN

Email: francesca.viani@bde.es

1 Introduction

Cross-border risk-sharing is a key dimension for assessing the performance of international business cycle models — see e.g. Obstfeld and Rogoff [2001] and Chari, McGrattan and Kehoe [2002]. The main focus of the literature is on the evidence emphasized by Backus and Smith [1993], who first calls attention on the fact that the correlation between the real exchange rate and relative consumption across countries tends to be low or even negative — at odds with the high degree of international risk-sharing predicted by standard open-economy models. Under complete markets and stable symmetric preferences, indeed, the correlation between relative consumption and real exchange rates must be close to one. A similarly high correlation turns out to be predicted also by many models assuming incomplete markets. In the last decade, a number of contributions have explored different mechanisms which can account for this evidence (typically referred to as the 'Backus-Smith puzzle'), ranging from strong asymmetric wealth effects in incomplete market economies (reflecting either terms of trade movements, or nontraded goods' prices) to marginal utility shifts (weakening the stark prediction of complete-market models regarding the connection between relative consumption and real depreciation) — see e.g. Benigno and Thoenissen [2006], Corsetti, Dedola and Leduc [2008], Ghironi and Melitz [2004], Mandelman, Rabanal, Rubio-Ramírez and Vilán, [2010], Nam and Wang [2010], Opazo [2006], Raffo [2010] among others.

In this paper, we decompose the Backus-Smith (henceforth BS) correlation in its dynamic components at high, business cycle, and lower frequencies. We make two contributions to the literature. First, we show that, among the OECD countries, the dynamic correlation between relative consumption and real depreciation tends to be low and negative at all frequencies, but more so at business-cycle and lower frequencies. In some cases, a low but positive overall correlation results from an average between a positive dynamic correlation at high frequencies, and a negative one at low frequencies. In this sense, spectral analysis underscores the point that contemporaneous correlations (especially when computed with first-differenced data) may give a somehow distorted picture, placing too much weight on the positive values at high frequencies — the BS puzzle is 'worse than you think'.

Our empirical analysis also shows that the (negative) dynamic correlation is more likely to be significantly different from zero at business cycle and lower frequencies, than at high frequencies. This finding casts doubts on the interpretation of the BS puzzle as a manifestation of the so-called 'exchange rate disconnect' puzzle. On the contrary, the puzzle is pervasive, and significantly so, at the appropriate frequencies for confronting the international business cycle models with the data, strengthening the case for placing international risk-sharing centerstage in the development of international economics.

Second, we provide a decomposition of the dynamic correlation predicted by standard open-economy models, into the sum of two terms. One term is constant across frequencies, and can be negative as a function of uninsurable risk. The second term is variable across frequencies, but necessarily positive, reflecting

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the insurance intertemporal trade provides against forecastable contingencies. Its strength over different frequencies will depend on the specific propagation mechanism embedded in a model.

In light of this analysis, in the last part of the paper we carry out an exercise applying spectral decomposition to simulated data from alternative open economy models proposed by the literature to match the BS correlation. These models typically feature asset-market imperfections, but also stress different international transmission mechanisms.

As is well understood, asset market imperfections *per se* are not enough for open economy models to generate significant deviations from the perfect correlation prediction. Incomplete markets must be complemented by transmission mechanisms which amplify the amount of uninsurable risk generated by business cycle impulses. A possible mechanism features a dominant role of the relative price of nontraded goods in driving real exchange rate movements and thus, when productivity shocks in the traded goods sector are also the prevailing source of fluctuations, Balassa-Samuelson effects. According to this mechanism, positive output gains in domestic tradables simultaneously drive up relative consumption and nontradable relative prices, while international tradable prices barely move or even fall (as in Benigno and Thoenissen (2006) and Devereux, Smith and Yetman (2009)). Another mechanism features endogenous income effects from output shocks to both tradables and nontradables, whereas, irrespective of the sectoral origin of the shock, output (productivity) increases cause all international relative prices of a country to strengthen, together with a rise in relative consumption (as in Corsetti, Dedola and Leduc (2008) or Ghironi and Melitz (2004)). We show that these mechanisms are able to deliver both an overall negative Backus-Smith correlation, and a negative dynamic correlation at business cycle frequencies.

The paper is organized as follows. Section 2 works out an analytical framework for analyzing the prediction by open-economy models, regarding the BS correlation at different frequencies. Section 3 carries out our empirical analysis. Using simulations, section 4 compares the performance of open economy models with the data. The appendices include a description of the data and the model, together with a robustness analysis of our empirical results.

2 The Backus-Smith evidence as a hurdle for open-economy model

Consider a standard open economy model with a domestic and a foreign country. In this framework, the equation pricing Arrow-Debreu bonds and the law of one price in the asset market imply that the growth of marginal utility of consumption, expressed in the same units, is equalized across countries state by state:

$$\beta \frac{U_C\left(C_t\right)}{U_C\left(C_{t-1}\right)} = \beta \frac{U_C^*\left(C_t^*\right)}{U_C^*\left(C_{t-1}^*\right)} \frac{RER_{t-1}}{RER_t} \tag{1}$$

where β denotes the discount rate (for simplicity assumed to be identical across borders), U_C and U_C^* denote the marginal utility of consumption, C and C^* denote consumption, in the domestic and the foreign economy respectively; RER is the real exchange rate, defined as the relative price of foreign consumption P^* in terms of domestic consumption P, i.e., $RER = P^*/P$. Under a symmetric specification of the two countries, perfect risk-sharing further implies that marginal utilities of consumption, again expressed in the same units, are equalized in levels, i.e.

$$U_{c,t} = U_{c^*,t}^* \frac{1}{RER_t} \tag{2}$$

Intuitively, the consumption allocation across countries should be such that the marginal benefit from an extra unit of domestic consumption equals its marginal cost, given by the foreign marginal utility of domestic consumption, times the relative price of C_t in terms of C_t^* (the inverse of the domestic real exchange rate). If a complete set of state-contingent securities is available, the above condition holds in a decentralized equilibrium independently of trade frictions and goods market imperfections (including shipping and trade costs, as well as sticky prices or wages), even when these frictions and imperfections cause deviations from the law of one price and failure of purchasing power parity (PPP).[1]

Under the assumption that, in each country, the national representative agent has preferences represented by a time-separable, constant-relative-risk-aversion utility function of the form $\dfrac{C^{1-\sigma} - 1}{1-\sigma}$, with $\sigma > 0$, the expressions (1) and (2) translate into conditions on the correlation between the (logarithm of the) ratio of domestic to foreign consumption and the (logarithm of the) real exchange rate RER,[2] respectively, in growth rates:

$$\frac{C_t^{-\sigma}}{C_{t-1}^{-\sigma}} = \frac{(C_t^*)^{-\sigma}}{(C_{t-1}^*)^{-\sigma}} \frac{RER_{t-1}}{RER_t}$$

and in level

$$RER_t = \left(\frac{C_t}{C_t^*}\right)^{\sigma}$$

Against the hypothesis of perfect risk-sharing, many empirical studies have found these correlations to be significantly below one, or even negative — in addition to the seminal paper by Backus and Smith [1993], see e.g. Kollmann [1995], and Ravn [2001] among others.

A key question addressed by the literature is under what conditions, if any, the empirical evidence can be reconciled with models of the international business cycle which do not assume complete markets. In models in which only an

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[1] It is only when PPP holds (i.e., $RER = 1$) that efficient risk-sharing implies equalization of the ex-post marginal utility of consumption. Only under this strong assumption, and ruling out shocks to preferences, complete markets imply a perfect cross-country correlation of consumption.

[2] Lewis [1996] rejects non-separability of preferences between consumption and leisure as an empirical explanation of the low correlation of consumption across countries.

uncontingent bond is traded across borders, for instance, (1) will not hold in general state by state, but only in expectations:

$$E_{t-1}\left[\beta\frac{C_t^{-\sigma}}{C_{t-1}^{-\sigma}}\right] = E_{t-1}\left[\beta\frac{(C_t^*)^{-\sigma}}{(C_{t-1}^*)^{-\sigma}}\frac{RER_{t-1}}{RER_t}\right] \tag{3}$$

When markets are incomplete, indeed, not all risk is insurable. Relative to the case of complete markets, the ex-post differential in expected utility growth (measured in the same unit) will be equalized up to an i.i.d. stochastic variable, itself a function of fundamental shocks (see e.g. Obstfeld 1994, Cochrane 2004), that is,

$$\frac{C_t^{-\sigma}}{C_{t-1}^{-\sigma}} + \zeta_t = \frac{(C_t^*)^{-\sigma}}{(C_{t-1}^*)^{-\sigma}}\frac{RER_{t-1}}{RER_t} \tag{4}$$

whereas $E_{t-1}(\zeta_t) = 0$. Uninsurable risk breaks, ex-post, the positive link between the growth rates of C/C^* and RER, which trade in a bond can ensure only ex-ante (in expectations).

It is because of uninsurable risk that standard models with incomplete markets may be able to predict an overall negative correlation between these variables, matching the BS evidence. But for being successful in this dimension, the amount of uninsurable risk in the model must be sufficiently large. First, it is well understood that, if shocks are purely transitory, restricting international trade to an international, uncontingent bond may cause the market allocation to be quite close to the complete-market one. Intuitively, when agents in one country get a temporary positive output shock, they will want to lend to the rest of the world, so that consumption increases both at home and abroad (see e.g., Baxter and Crucini [1995]). In addition, the international transmission of shocks is also shaped by relative price movements. If higher output is associated, in equilibrium, to lower international prices of domestic goods, higher output in one country benefits foreign consumers by boosting their income in real terms — a mechanism which automatically contributes to production risk-sharing. Under some restrictions on the model's parameters, relative price movements can actually ensure complete sharing of production risk, independently of trade in financial assets — a point underscored by Cole and Obstfeld [1991], Corsetti and Pesenti [2001, 2005] and Corsetti, Dedola and Leduc [2010].

These considerations suggest that otherwise standard open-economy models with imperfect asset markets (and stable preferences) may predict a low or negative correlation between relative consumption and real depreciation when (a) shocks are persistent and (b) the core transmission mechanism mutes, or reverses, the role of relative price adjustment in providing risk-sharing. Several contributions in the literature exploit this very insight, to identify international transmission mechanisms through which business cycle disturbances translate into large ex-post wedges between marginal utilities. As discussed in the introduction, some contributions in the literature emphasize endogenous income effects from output fluctuations which, irrespective of their sectoral origin, cause all international relative prices of a country, but especially tradable prices, to

co-move inversely with relative output and relative consumption (as in Corsetti, Dedola and Leduc [2008] or Ghironi and Melitz [2004]). Other contributions emphasize a different mechanism, hinging upon a dominant role of the relative price of nontraded goods in driving real exchange rate movements (as in Benigno and Thoenissen [2006], or Cova, Pisani, Batini and Rebucci, [2008]).

Without loss of generality, up to a first order of approximation the solution for the differential in consumption growth and the exchange rate growth in these models can always be written in terms of their state-space representation:

$$\sigma \Delta(\widehat{C}_t - \widehat{C}_t^*) = \pi_1^C \mathbf{S}_{t-1} + \pi_2^C \varepsilon_t \equiv x$$
$$\Delta\left(\widehat{RER}_t\right) = \pi_1^R \mathbf{S}_{t-1} + \pi_2^R \varepsilon_t \equiv y$$

where S_{t-1} is the vector of state variables (endogenous and exogenous) pre-determined at t, including lagged variables, and thus orthogonal to the vector of fundamental $i.i.d.$ shocks ε_t.[3] Using these expressions, we can provide a crucial insight on the properties of the correlation between our variables of interest. Specifically, the unconditional correlation can be readily derived from the covariance and variances below:

$$Cov(x_t, y_t) = \pi_1^C Var(\mathbf{S}_{t-1})\pi_1^{R\prime} + \pi_2^C Var(\varepsilon_t)\pi_2^{R\prime} \quad (5)$$
$$Var(x_t) = \pi_1^C Var(\mathbf{S}_{t-1})\pi_1^{C\prime} + \pi_2^C Var(\varepsilon_t)\pi_2^{C\prime}$$
$$Var(y_t) = \pi_1^R Var(\mathbf{S}_{t-1})\pi_1^{R\prime} + \pi_2^{R\prime} Var(\varepsilon_t)\pi_2^{R\prime}$$

In the case of complete markets, the ex-post equalization of the RER-weighted growth rates in marginal utilities (under the simplifying assumption of CRRA and separable preferences, i.e., $\sigma \Delta(\widehat{C}_t - \widehat{C}_t^*) = \Delta\left(\widehat{RER}_t\right)$), implies

$$\pi_1^C = \pi_1^R = \pi_1 \quad (6)$$
$$\pi_2^C = \pi_2^R = \pi_2$$

It is then straightforward to verify that, holding these restrictions, the covariance in (5) will be the sum of two quadratic expressions, hence always positive. Moreover, it will be identically equal to each of the variances,

$$Cov(x_t, y_t) = \pi_1 Var(\mathbf{S}_{t-1})\pi_1' + \pi_2 Var(\varepsilon_t)\pi_2' = Var(x_t) = Var(y_t)$$

ensuring that the BS correlation coefficient will be equal to one.

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[3]It is clear that if the variable $\widehat{X} = \widehat{C}_t, \widehat{C}_t^*$, and $\widehat{RER}_t$ in levels, have a state-space representation

$$\widehat{X}_t = \nu \mathbf{Z}_{t-1} + \eta \varepsilon_t,$$

then their growth rates will have the representation assumed in the text with

$$\mathbf{S}_{t-1} = \left[\begin{array}{c} \mathbf{Z}_{t-1} \\ \widehat{\mathbf{X}}_{t-1} \end{array} \right]$$

and corresponding coefficient matrices.

The restrictions (6) do not necessarily hold under incomplete markets, however. Namely, in a bond economy, the equalization of the growth rate of the RER-weigthed marginal utilities in expectations:

$$\begin{aligned} E_{t-1}\sigma\Delta(\widehat{C}_t - \widehat{C}_t^*) &= \pi_1^C \mathbf{S}_{t-1} = \\ E_t\Delta\left(RER_t\right) &= \pi_1^{RER}\mathbf{S}_{t-1}, \end{aligned}$$

only implies

$$\pi_1^C = \pi_1^R = \pi_1$$

without restricting π_2^C and π_2^R to be identical. Rewriting the covariance and variances under the appropriate restriction,

$$\begin{aligned} Cov(x_t, y_t) &= \pi_1 Var(\mathbf{S}_{t-1})\pi_1' + \pi_2^C Var\left(\varepsilon_t\right)\pi_2^{R\prime} \\ Var\left(x_t\right) &= \pi_1 Var(\mathbf{S}_{t-1})\pi_1' + \pi_2^C Var\left(\varepsilon_t\right)\pi_2^{C\prime} \\ Var\left(y_t\right) &= \pi_1 Var(\mathbf{S}_{t-1})\pi_1' + \pi_2^{R\prime} Var\left(\varepsilon_t\right)\pi_2^{R\prime} \end{aligned}$$

makes clear that, in a bond economy, only the first term of the covariance is a quadratic form, always positive. As discussed further below, this term is related to the insurance intertemporal trade provides against predictable contingencies.

The second term can instead be negative, depending on the properties of the model concerning the uninsurable risk ζ_t. In the models by Corsetti, Dedola and Leduc [2008], Ghironi and Melitz [2004] and Benigno and Thoenisson [2006], for instance, relative price movements magnify the uninsurable component of fundamental risk, up to moving the real exchange rate and the consumption differential in opposite directions. In other words, the term $\pi_2^C Var\left(\varepsilon_t\right)\pi_2^{R\prime}$ is not only negative, but actually larger, in absolute value, than the quadratic form in the variance of S.

For a bond economy, the above expressions have crucial implications for the analysis in the frequency domain. Recall that the spectrum of i.i.d. vectors is constant across frequencies (see e.g. Hamilton 1994). So, whether or not the second term in the covariance above is negative and large in the model, its value will nonetheless be constant across the spectrum. Conversely, the spectrum of the quadratic form in the variance of the state vector will generally not be constant, reflecting the fact that S can also be represented as a VAR. It follows that in a bond economy the strength and even the sign of the correlation can be expected to vary across frequencies, depending on the dynamics of the variance of the vector of state variables, $Var(S_t)$.

A key conclusion is that, in standard incomplete-market models, differences across frequencies in the cospectrum of the BS correlation are driven by the insurable component in $\Delta(\widehat{C}_t - \widehat{C}_t^*)$ (and/or $\Delta\left(RER_t\right)$) — a component reflecting the endogenous dynamics of the state variables and thus the specific features of the propagation mechanism embedded in each model. The properties of the dynamic Backus-Smith correlation at different frequencies thus map into the theoretical predictability of the rate of real depreciation (or the differential in consumption growth). For instance, if in equilibrium the real exchange rate

follows a random walk — a common view in the literature — implying:

$$E_t \Delta (RER_t) = \pi_1 \mathbf{S}_{t-1} = 0,$$

then in a bond economy the dynamic correlation should be expected to be constant across all frequencies. A random walk for the exchange rate by no means implies that the exchange rate is 'disconnected' from the fundamentals. As shown above, the contemporaneous covariance with the cross-country differential in the growth rate of consumption will generally be non-zero, even in the case of a bond economy, because of uninsured risk. However, the fact that the model predicts no forecastability for the rate of depreciation has strong implications for the amount of insurance provided by trade in bonds. Indeed, it can be shown that in the random walk case the bond will not be use in equilibrium.[4]

To appreciate the implications of these considerations for our analysis, consider the simplest possible setup, with one state and one variable. The covariance can be written

$$Cov(x,y) = \left(\frac{\pi_1^2 \mu^2}{1 - \rho^2} + \pi_2^C \pi_2^{RER} \right) Var(\varepsilon)$$

implying that the cospectrum varies across frequencies according to the following

$$\left[\frac{\pi_1^2 \mu^2}{1 + \rho^2 - 2\rho Cos(\omega)} + \pi_2^C \pi_2^{RER} \right] \frac{var(\varepsilon)}{2\pi}.$$

In this example, for the covariance to be negative, it must be the case that:

$$\frac{\pi_1^2}{1 - \rho^2} < -\frac{\pi_2^C \pi_2^{RER}}{\mu^2}.$$

In the univariate case, then, a negative covariance necessarily implies a negative cospectrum at any frequency if $\pi_1^2 = 0$ (the random walk case) or more generally if:

$$\frac{1}{1 - \rho^2} \quad \geq \quad \frac{1}{1 + \rho^2 - 2\rho Cos(\omega)}$$

$$\Longleftrightarrow \quad 2\rho \frac{\rho - Cos(\omega)}{1 + \rho^2 - 2\rho Cos(\omega)} \geq 0$$

$$\Longleftrightarrow \quad \rho \geq Cos(\omega).$$

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[4]Moreover, under rational expectations, trade in uncontingent bonds implies precise restrictions on the extent to which the differentials in consumption growth rates and the real exchange rate are forecastable. If in equilibrium, the real exchange rate follows a random walk, it must be the case that the consumption differential is also not forecastable:

$$E_t \Delta (RER_t) = \pi_1 \mathbf{S}_{t-1} = 0$$
$$<=>$$
$$E_{t-1} \sigma \Delta (\widehat{C}_t - \widehat{C}_t^*) = \pi_1^C \mathbf{S}_{t-1} = 0.$$

It is apparent that, for a stationary process with $|\rho| < 1$, such a condition would be impossible to satisfy at low frequencies, for $\omega \to 0$. Moreover, depending on whether ρ is above or below zero, the positive contribution $\dfrac{\pi_1^2 \mu^2}{1 + \rho^2 - 2\rho Cos(\omega)}$ to the cospectrum will become more relevant at lower or higher frequencies.[5]

Multivariate examples are much more complex to treat analytically, precluding the derivation of general conditions on the sign of the cospectrum. Yet, the main message from our simple example would survive. In general, an overall negative covariance between two variables does not necessarily imply a negative cospectrum at all frequencies. The question is therefore whether the one-bond (incomplete-market) economies addressing the BS puzzle in terms of unconditional correlations, would instead predict a positive dynamic correlation at some lower frequency, and whether this would be consistent with the data. The question boils down to verify whether the condition (3) implied by trade in bonds, imposes any restriction on the cospectrum in model economies. Intuitively, what is at stake is whether the amount of insurance provided by trade in bonds becomes so large at some frequency, that it results in a switch in the sign of the correlation over the spectrum.

3 A spectral analysis of the Backus-Smith correlation

In this section, we reconsider the BS evidence using spectral analysis, focusing on a sample of OECD countries. Namely, we estimate a measure of the contributions of cycles of different frequency to the Backus-Smith correlation.

As is well known (see e.g. Hamilton, Chapter 10.4), given any two covariance-stationary series x and y, the cospectrum is the frequency-domain equivalent of the covariance between them. Specifically, the cospectrum measures the portion of the covariance between x and y that is attributable to cycles of a given frequency ω. The correlation between relative consumption and real exchange rate at frequency ω is then measured by the dynamic correlation (see Croux et al. (2001))

$$\rho_{C,RER}(\omega) = \frac{C_{C,RER}(\omega)}{\sqrt{S_C(\omega) \cdot S_{RER}(\omega)}},$$

where $C_{C,RER}$ denotes the cospectrum between the two series and S their spectra at frequency ω.

Our sample consists of 20 OECD countries, for which we have quarterly data over the period 1971:1 and 2009:2. For each country, as customary in

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[5] Of course, the condition for the the cospectrum to be negative at any frequency is

$$\left| \frac{1}{1 + \rho^2 - 2\rho Cos(\omega)} \right| < -\frac{\pi_2^C \pi_2^{RER}}{\pi_1^2 \mu^2},$$

In general, conditions like the above cannot be assessed without the knowledge of the structural parameters of the model.

the literature, we study the correlation between the ratio of domestic to foreign consumption and the real exchange rate both vis-à-vis the US, as well as vis-à-vis a trade-weigthed aggregate of the other countries in the sample — an aggregate dubbed 'Rest of the World', or ROW. Results are shown in the panels A and B of Table 1, respectively. In each panel, the first two columns report the standard Backus-Smith statistics, i.e. the correlations for first-differenced data and HP-filtered data series. The last three columns instead display the dynamic correlations for first-differenced series. Business-cycle frequencies refer to a time horizon of 8 to 32 quarters, low frequencies to more than 32 quarters, and high frequencies to less than 8 quarters — technical details on the estimation are given in the appendix.

The first two columns in Table 1A reproduce the well-known result, that the BS correlation for the OECD countries against the US is low or negative, whether this is measured using differenced data or HP-filtered data. The evidence in the table is clearly in line with earlier studies: nearly all entries in the first two columns have a negative sign — this is true for 17 out of 19 cases when data are first-differenced, and 16 out of 19 cases when data are HP-filtered; even when positive, the correlation is typically close to zero.

Spectral analysis unveils an important new dimension of the evidence. The dynamic correlation is not constant, but remains mostly negative at all frequencies. At low and business-cycle frequencies, however, the inverse correlation tends to be stronger. At business-cycle frequencies, the number of countries with a negative correlation is 17 out of 19 countries, while all entries have a negative sign in the column corresponding to low frequencies. Conversely, the number of negative entries falls to 14 out of 19 at high frequencies. As regards the intensity of the inverse correlation, the average across countries goes from -.04 for high frequencies, to -.18 and -.22 for business-cycle and low frequencies, respectively. The highest values for the inverse correlation are recorded for Korea (-.55) at business-cycle frequencies, and Spain (-.53), Finland (-.47), Korea (-.48), and Sweden (-.47) at low frequencies. Also, note that the two countries showing an overall positive correlation (Austria and Japan), display a negative correlation at business-cycle and/or low frequencies. By way of example, the values for Austria are 0.4, -.11, and -.16, at high, business-cycle and low frequencies, respectively. This should not come as a surprise, since it is well known that standard correlations in the time domain with differenced data tend to boost the high frequency components of time series (see e.g. Croux et al. 2001).

Because of the different economic structure between the US and aggregates of OECD countries (spanning features from economic size and openness to policy regimes), it is reasonable to expect some variability in results when countries are assessed against the ROW. This is indeed the case in Table 1B. Comparing the first two columns of Table 1A and Table 1B, reproducing the standard statistics, the number of negative entries falls from 17 (out of 19) to 11 (out of 20). Even so the correlation never exceeds +.20 (for the first-differenced series). The minimum value is -.41, again for Korea.

Also in Table 1B, the Backus-Smith puzzle is more pronounced at business-cycle and low frequencies, whereas a negative correlation is recorded for 13 and

15 countries, respectively, against 8 countries at high frequencies (out of the 20 in the sample). The sample average ranges from -.16 (low frequencies) to .05 (high frequencies). Moreover, for four countries (Austria, Japan, Norway and New Zealand), a small positive value for the overall statistic in the first column of the table appears to be an average of negative and positive values at different frequencies.

A sharper picture emerges from the graphs displayed in Figure 1A and 2B, which plot the result of spectral analysis by frequency, together with a 5% and 95% confidence bands around the point estimates. These two figures, plotting the dynamic correlation for each country vis-à-vis the US and the ROW, respectively, emphasize a further key result. Not only the dynamic correlation is always significantly lower than unity — it is actually significantly lower than .5 at all frequencies in both figures. At business-cycle or lower frequencies, it is also significantly different from zero for 12 out of 19 countries in Figure 1A, and 14 out of 20 countries in Figure 1B. In Figure 1A, the statistically significant correlation has always a negative sign. In Figure 1B, there are four countries (Belgium, Switzerland Ireland and marginally Japan) which display a significantly positive correlation, consistent with Table 1B.

The importance of this result is best appreciated in light of a recent contribution by Rubio-Ramirez and Rabanal (2010), showing that the business-cycle and lower frequencies account for the bulk of the variability of the real exchange rate. Our evidence further suggests that it is precisely at the frequencies at which the real exchange rate is more volatile (but arguably also more predictable), that it is also significantly 'connected' to macroeconomic variables, like relative consumption.

Robustness is discussed in the appendix. While the use of first-differenced data is natural in the context of bond economies, empirically, it might boost the high frequency component of a series. For this reason, we re-do our analysis applying the bandpass filter described in Christiano and Fitzgerald (2003). The main conclusions from this robustness exercise are in line with the ones discussed above.

TABLE 1A

Correlation between real exchange rate and relative consumption vis-à-vis the US

| COUNTRY | Differenced | HP-filtered | Dynamic Correlation | | |
			Low freq	BC freq	High freq
Australia	-0,07	-0,24	-0,33	-0,10	-0,01
Austria	0,00	0,10	-0,16	-0,11	0,04
Belgium	-0,08	0,03	-0,20	-0,09	-0,04
Canada	-0,16	-0,10	-0,21	-0,24	-0,12
Switzerland	-0,06	-0,04	-0,12	0,02	-0,09
Denmark	-0,20	-0,26	-0,30	-0,25	-0,16
Spain	-0,21	-0,27	-0,53	-0,37	0,01
Finland	-0,10	-0,41	-0,47	-0,33	0,05
France	-0,07	-0,03	-0,22	-0,06	-0,04
Germany	-0,12	-0,06	-0,21	-0,18	-0,10
Ireland	-0,05	-0,18	-0,16	-0,01	-0,04
Italy	-0,01	-0,23	-0,25	-0,10	0,10
Japan	0,02	0,25	-0,14	0,01	0,05
Korea	-0,46	-0,48	-0,47	-0,55	-0,44
Netherlands	-0,12	-0,09	-0,22	-0,22	-0,08
Norway	-0,06	-0,16	-0,13	-0,13	-0,03
New Zealand	-0,04	-0,36	-0,44	-0,27	0,10
Sweden	-0,17	-0,30	-0,45	-0,20	-0,10
UK	-0,08	-0,41	-0,42	-0,34	0,02
US	NA	NA	NA	NA	NA
Median	-0,08	-0,18	-0,22	-0,18	-0,04

NOTE: See the data appendix for a description of the data used.

TABLE 1B

Correlation between real exchange rate and relative consumption vis-à-vis the rest-of-the-world

COUNTRY	Differenced	HP-filtered	Dynamic Correlation		
			Low freq	BC freq	High freq
Australia	-0,05	-0,19	-0,32	-0,01	-0,03
Austria	0,07	0,00	-0,27	-0,14	0,12
Belgium	0,12	0,51	0,21	0,25	0,04
Canada	-0,13	-0,05	-0,12	-0,21	-0,10
Switzerland	0,19	0,16	0,29	0,26	0,15
Denmark	-0,06	-0,17	-0,15	-0,05	-0,05
Spain	-0,05	-0,17	-0,35	-0,21	0,11
Finland	-0,07	-0,45	-0,55	-0,36	0,14
France	0,11	0,21	0,13	0,12	0,10
Germany	-0,02	0,07	-0,07	0,10	-0,05
Ireland	0,19	0,30	0,32	0,32	0,12
Italy	-0,08	-0,27	-0,32	-0,12	-0,03
Japan	0,09	0,32	-0,07	0,12	0,10
Korea	-0,41	-0,51	-0,49	-0,53	-0,38
Netherlands	-0,04	-0,07	-0,17	-0,18	0,01
Norway	0,03	0,00	-0,01	-0,04	0,05
New Zealand	0,05	-0,32	-0,44	-0,17	0,16
Sweden	-0,15	-0,23	-0,28	-0,24	-0,10
UK	0,06	0,07	0,02	0,02	0,09
US	-0,20	-0,15	-0,36	-0,26	-0,14
Median	-0,03	-0,06	-0,16	-0,09	0,05

NOTE: The rest-of-the-world is a trade-weighted aggregate of all the countries in the same table. See the data appendix for a description of the weights.
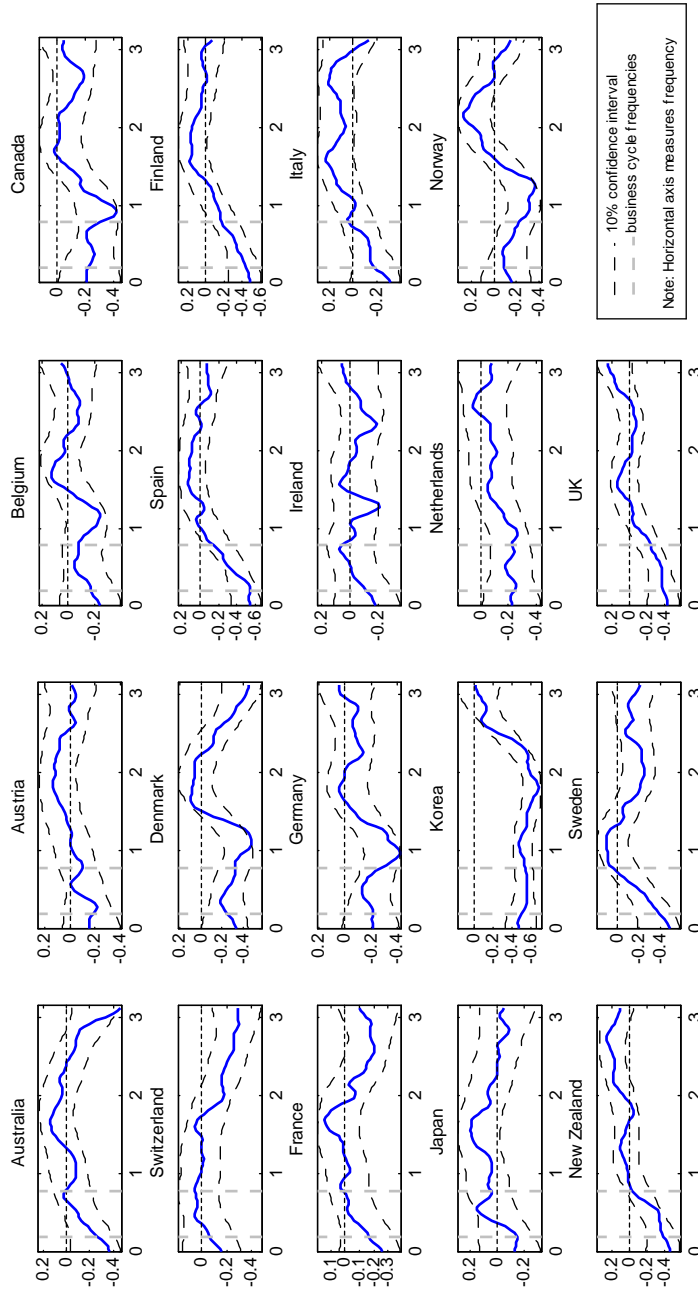
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FIGURE 1A − Dynamic correlations of relative consumption and real exchange rates vis-à-vis the US.
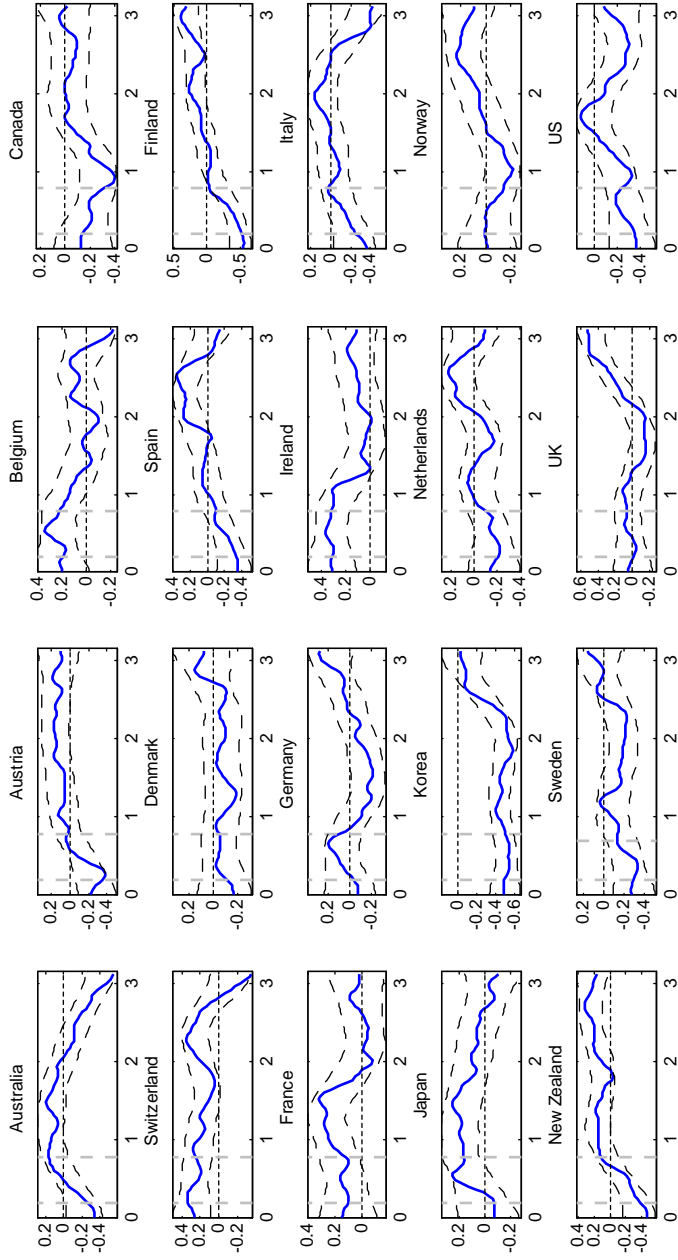
FIGURE 1B − Dynamic correlations of relative consumption and real exchange rates *vis-à-vis* the Rest of the World.

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4 The 'hurdle' in frequency domain

In the previous section, we have cast new light on the evidence at odds with open-economy models predicting a high degree of international risk insurance. The novel contribution of our analysis is that, in the data, the BS puzzle is pervasive at business-cycle and lower frequencies, more so than at higher frequencies. This section is devoted to spell out some implications of this evidence for open economy modelling.

To start with, our results clarify that a low value of the BS statistics should not be mistaken for a manifestation of the so-called exchange rate 'disconnect puzzle.' For many countries, indeed, once high frequencies are eliminated, the BS correlation at business-cycle and lower frequencies is significantly different from zero at standard levels. By the same token, our results make the notion that this evidence provides a meaningful hurdle for open-economy models, much more compelling. If the BS correlation were to be driven by correlations at high frequencies, for instance, one could argue that market dynamics at high frequencies are not appropriately captured by models designed to account for macro dynamics at business-cycle and lower frequencies. On the contrary, our evidence makes it clear that the BS puzzle is pervasive exactly at the frequencies over which open economy models appear to be relatively successful in matching the data in a number of dimensions. Also in this respect, our analysis strengthens the case for placing the analysis of international risk-sharing centerstage in the development of international economics, as advocated e.g. by Obstfeld and Rogoff [2001] and Chari et al. [2002].

In the literature, different contributions propose different mechanisms by which the transmission of shocks induce a negative correlation between relative consumption and real depreciation. In principle, these mechanisms can create a variety of state dynamics, which may or may not fit our evidence from spectral analysis. In what follows, we set up a two-country model with tradables and nontradables, and simulate it under different calibrations of the shock processes and parameters specification, so as to make it broadly consistent with the key quantitative results in Corsetti et al. [2008], and Benigno and Thoenissen [2006] in terms of the Backus-Smith correlation. The first paper, by Corsetti et al. [2008], actually proposes two distinct transmission mechanisms, one based on a relatively low trade elasticity, the other on a high trade elasticity and persistent shocks. In either specification, what generates a negative unconditional correlation is a strong wealth effect from shocks, translating into a negative correlation between relative consumption and the terms of trade. Overall, we will thus consider three different specifications of the model. Since most features of the analysis are standard, we include a description of the general structure of model in the appendix. Note that, in our exercises, productivity shocks are both country- and sector-specific, Define $Z = [A_T, A_T^\star, A_{NT}, A_{NT}^\star]$, where A denotes TFP and the subscript T and NT refers to tradables and nontradables, respectively. We consider the process

$$Z\prime = \lambda Z + u,$$

where technology shocks are identified with Solow residuals in each sector.

All the models under consideration are calibrated to the US versus the rest of the OECD countries, with annual data. Since the dynamic correlations for the US in Table 1B and Figure 1B are computed using quarterly data, the relevant empirical comparison is carried at comparable frequencies by properly adjusting the parameter ω. For instance, business cycle frequencies are still assumed to include cycles from 2 to 8 years, namely for $\omega = [\pi/2, \pi/8]$.

Note that, for the US, the BS correlation tends to be significantly negative at business-cycle and lower frequencies. Moreover, the 90% confidence bands shown in Figure 1B suggest that the value of this correlation is likely to be rather flat at these frequencies. We focus on the following question: can these bond-economy models accounting for a negative BS correlation also account for a negative dynamic correlation at the relevant frequencies of the spectrum?

In the first specification under consideration, we calibrate the model corresponding to the low-trade-elasticity economy in Corsetti et al. [2008]. We set the elasticity of substitution between domestic and foreign tradables equal to 0.74.[6] Because of distribution costs, the trade elasticity is half the elasticity of substitution, thus equal to .375. Productivity shocks are from Corsetti et al. [2008], and estimated using annual data in manufacturing and services from the OECD STAN database.

In the second specification, we calibrate the model as in the high-elasticity high-shock-persistence case discussed in the same paper. The trade elasticity is raised to a value as high as 4, consistent with the estimates by Bernard et al. [2003]. With a high elasticity, strong wealth effects require high persistence of the shocks: the autoregressive coefficient for the shocks hitting the tradable good sector is set close to one — while spillovers are set to zero to guarantee stationarity.[7]

The third model specification aims at reproducing the results in, and the transmission mechanism envisioned by, Benigno and Thoenissen [2006]. These authors set the trade elasticity equal to 2, and estimate the productivity process using a different sample and dataset (from the Groningen Growth and Development Centre). Because of these differences, the estimates of the shock process is quite different relative to the one in CDL — the most apparent difference referring to the persistence of the shock in the tradable and nontradable sector, whose ranking is reversed (with the shocks to nontradables much less persistent). For the BT mechanism to be effective in addressing the BS puzzle, indeed, the shocks to the nontradable sector cannot be prominent. Output (productivity) disturbances in nontradables indeed tends to mute the mechanism working through nontradable prices since, other things equal, these disturbances would drive relative consumption and the real exchange rate in the same direction: the unconditional Backus-Smith correlation could then have either sign.

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[6] This value is slightly lower than the one used in the benchmark bond economy in CDL, depending on the specification of investment.

[7] The role of shock persistence is also explored by Opazo [2006] and Nam and Wang [2010], who consider news shocks, as well as by Corsetti, Dedola and Leduc [2010], also in relation to the design of optimal policy in open economies.

Starting with a model whose structure is marginally different from the one adopted by Benigno and Thoenissen [2006], we thus proceed by adjusting the CDL calibration as to reproduce the BT results. Specifically, in the baseline shock process estimation by CDL, we increase the autocorrelation of tradable shocks, and decrease the standard deviation of nontraded shocks. Moreover, we set the share of labour in tradables equal to 0.33 (and the trade elasticity equal to 2) as in Benigno and Thoenissen (2006). With these adjustments, the model predicts a negative Backus-Smith correlation driven by movements in the relative price of nontraded goods. As a caveat, however, we should stress that results are sensitive not only to the relative weight of shocks to nontradables as drivers of the business cycle, but also to increasing the trade elasticity, as to approximate the standard textbook Balassa-Samuelson model. Both the shock process and the value of the trade elasticity turn out to be key for the model to match the BS correlation.

The outcome of our exercises are summarized in Table 2, showing the unconditional correlation and the dynamic correlation arising at business cycle and low frequencies, both computed using first-differenced data. In the Table, CDL refers to the two specifications after Corsetti et al. [2008], BT to that from Benigno and Thoenissen [2006]. For the CDL models, we also report results assuming nominal rigidities, with an average duration of prices of 4.3 months (matching the evidence in Bils and Klenow [2004]). We instead omit the simulations of the BT specification with nominal rigidities, as in this case results appear to be quite sensitive also to price stickiness — the BS correlation turns positive and high (although not perfect) across the spectrum.

We include in the table results from the standard specification of the model, for an elasticity equal to 1.5. While providing the motivation for the discussion of the BS puzzle in the first place, these results are also helpful in reconciling the models we use in our simulations, assuming non-separable preferences in consumption and leisure as typical in the literature, with the analysis in Section 2, where we have assumed separability for clarity of exposition. The fact that, with an elasticity of 1.5, the BS correlation is close to one at all frequencies, suggests that non-separability play no meaningful role in our experiments. The last line reproduces the relevant empirical results from Table 2B, together with the extremes of the 5% to 95% confidence band.

It is apparent that all the models under consideration (at least with flexible prices) are able to generate low and even negative dynamic correlations at business-cycle and lower frequencies. For a variety of transmission mechanisms amplifying the amount of uninsurable risk in the economy due to productivity shocks, trade in one bond does not seem to provide an effective instrument to insure risks at any frequency. In some cases our simulations produce correlations well within the confidence band of Figure 1B. For instance, this is the case of the CDL specification with a low elasticity (first two rows of Table 2), which overall comes reasonably close to the evidence for the US. In our exercises, the CDL specification with high elasticity and persistent shocks produces a correlation which is too negative at business cycle frequencies, relative to the evidence. At the same frequencies, the BT specification errs on the other side — in our

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specification the BT model is also excessively sensitive to nominal rigidities.

TABLE 2

Correlation between real exchange rate and relative consumption
in the simulated model and in the data

Model specification	Differenced	Dynamic Correlation	
		Low frequency	BC frequency
CDL low elasticity	-0.14	-0.20	-0.13
With nominal ridigities	-0.2	-0.14	-0.22
CDL high elasticity	-0.19	-0.20	-0.62
With nominal rigidities	-0.78	-0.60	-0.89
BT	-0.15	-0.51	-0.06
Standard model	0.99	0.99	1
Data (US vs ROW)	-0.20	-0.36	-0.26
Confidence int. Upper bound		-0.121	-0.125
Confidence int. Lower bound		-0.524	-0.522

NOTE: CDL: Corsetti, Dedola and Leduc (2008); BT: Benigno and Thoenissen (2006); low elasticity = 0.74; high elasticity = 8. Nominal rigidities: average duration of price-stickiness = 4.3 months (from Bils and Klenow [2004]). Correlations computed over 5000 simulations. Frequency bands are defined as in Table 1A. Confidence intervals on US data are computed over 500 bootstrap simulations.

5 Conclusions

Over two decades the Backus-Smith "puzzle" has challenged models of the international transmission. In this paper, we have provided theoretical and empirical arguments for redefining the puzzle. First, by using spectral analysis, we have shown that the BS evidence is actually much starker at business cycle and lower frequencies, than suggested by contemporaneous correlation. In the data, the correlation between relative consumption and the real exchange rate for many countries is significantly negative, exactly at those frequencies which are most appropriate for assessing the performance of international business cycle models.

Second, we have shown that, in standard open-economy models, the prediction of a negative unconditional correlation does not necessarily imply a negative sign for the cospectrum. Dynamic correlations can instead be expected to vary as a function of structure of the model economy, determining the amount of

insurance provided by trade in the assets available to economic agents. With incomplete markets, it may well be possible that a model predicting a negative contemporaneous correlation, fails to match the evidence over the spectrum, if trade in bonds provides significant insurance against predictable contingencies at some relevant frequency. Leading models in the literature, exploring mechanisms which can account for a low degree of international risk-sharing, generate predictions that appear to be broadly in line with the evidence.

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A   Appendix. Data sources

We collected quarterly data on real consumption from the OECD Economic Outlook for Australia, Austria, Belgium, Canada, Switzerland, Denmark, Spain, Finland, France, Germany, Ireland, Italy, Japan, Korea, the Netherlands, Norway, New Zealand, Sweden, the UK and the US. Consumer price indexes and nominal exchange rates are quarterly data from the IMF International Financial Statistics database for the period 1971:1-2009:2.

For each country in our dataset, the Foreign counterpart is either the US or a trade-weighted aggregate of all the other countries in the sample. In the latter case, S_t and $P_t^\star$ are trade-weighted averages of all the other countries in the dataset. Trade weights were built computing bilateral trade shares. Namely, we computed the trade share of country i from country j as

$$0.5 \cdot \frac{\exp_j^i}{\exp^i} + 0.5 \cdot \frac{\mathrm{imp}_j^i}{\mathrm{imp}^i}$$

where $\exp_j^i$ and imp_j^i are exports and imports from country j, and $\exp^i$ and imp^i denote total exports and imports of country i. Exports and imports are averages of annual data over the period 1980-2008, collected from the IMF Direction of Trade Statistics. For the median country trade weights account for roughly 73% of total imports and exports.

B   Appendix. Spectral analysis

Spectra and cospectra are estimated non-parametrically using a smoothing window of length $m = (T)^{1/2}$, where T is the sample size. In particular, we use a Tukey window, which associates any linearly-spaced vector x to

$$w\left(x\right) = \left\{ \begin{array}{ll} \frac{1}{2} \cdot \left\{1 + \cos\left(\frac{2\pi}{r}\left[x - r/2\right]\right)\right\}, & 0 \leq x \leq r/2 \\ 1, & r/2 \leq x \leq 1 - r/2 \\ \frac{1}{2} \cdot \left\{1 + \cos\left(\frac{2\pi}{r}\left[x - 1 + r/2\right]\right)\right\}, & 1 - r/2 \leq x \leq 1 \end{array} \right\},$$

where r is the smoothing parameter indicating the ratio of taper to constant section in the window, and is assumed to be equal to 0.5.[8]

We build confidence intervals from 500 bootstrap replicates. For the dynamic correlation between relative consumption and real exchange rates, we use sigma-confidence intervals. More specifically, we apply the Fisher-z transformation to the simulated dynamic correlations in order for their distribution to get closer to a normal, compute sigma-intervals on the transformed series, and finally convert them into bands for the dynamic correlation.[9]

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[8] The Tukey window collapses to a rectangular window for $r = 0$ and to a Hanning window for $r = 1$. Results obtained with these two alternative parametrizations are available upon request.

[9] See Croux et al. (2001).

C   Appendix. Bandpass filter approach

Since first-differencing might boost the low frequency component of a series, we check the robustness of our results using an alternative approach. Namely, we apply to the log-levels of relative consumption and real exchange rate, the bandpass filter described in Christiano and Fitzgerald (2003), and compute the correlation arising at different frequency bands. We used the following bands. High frequency corresponds to cycles of length between 2 and 5 quarters, business-cycle frequency to cycles between 6 and 32 quarters, low frequency to cycles between 33 and 70 quarters. The following tables report the main results.

TABLE 3A

Correlation between real exchange rate and relative consumption vis-à-vis the US
Bandpass-filtered data

| COUNTRY | Correlation | | |
	Low frequency	BC frequency	High frequency
Australia	-0,14	-0,41	-0,04
Austria	-0,20	-0,10	0,04
Belgium	-0,29	-0,23	-0,10
Canada	-0,72	0,06	-0,10
Switzerland	0,35	-0,35	-0,14
Denmark	-0,41	-0,10	-0,23
Spain	-0,84	-0,32	-0,04
Finland	-0,14	-0,60	0,00
France	-0,33	-0,22	-0,06
Germany	-0,26	-0,22	-0,17
Ireland	-0,61	-0,07	-0,12
Italy	-0,20	-0,33	0,08
Japan	0,01	0,20	0,06
Korea	-0,71	-0,58	-0,46
Netherlands	-0,39	-0,19	-0,09
Norway	-0,01	-0,16	-0,15
New Zealand	-0,19	-0,57	0,12
Sweden	-0,74	-0,33	-0,08
UK	-0,34	-0,36	0,04
US	NA	NA	NA
Median	-0,29	-0,23	-0,08

 NOTE: See the data appendix for a description of the data used
High freq: 2-5 quarters; business cycle freq: 6-32 quarters; low freq: 33-70 quarters

TABLE 3B

Correlation between real exchange rate and relative consumption vis-à-vis the
rest-of-the-world – Bandpass-filtered data

COUNTRY		Correlation	
	Low frequency	BC frequency	High frequency
Australia	-0,19	-0,27	0,00
Austria	-0,31	-0,01	0,12
Belgium	0,03	0,34	0,09
Canada	-0,59	0,10	-0,10
Switzerland	0,76	-0,03	0,16
Denmark	-0,28	0,16	-0,12
Spain	-0,79	-0,03	0,11
Finland	-0,59	-0,64	0,05
France	0,07	0,15	0,15
Germany	-0,07	0,15	-0,10
Ireland	0,35	0,37	0,04
Italy	-0,63	-0,17	-0,03
Japan	0,57	0,31	0,13
Korea	-0,69	-0,60	-0,41
Netherlands	-0,23	0,06	0,06
Norway	0,01	0,18	-0,07
New Zealand	-0,41	-0,49	0,17
Sweden	-0,59	-0,20	-0,11
UK	-0,14	0,32	0,11
US	-0.66	-0,21	-0,14
Median	-0,25	0,03	0,04

 NOTE: The rest-of-the-world is a trade-weighted aggregate of all the countries in the
same table. See the data appendix for a description of the weights
High freq: 2-5 quarters; business cycle freq: 6-32 quarters; low freq: 33-70 quarters
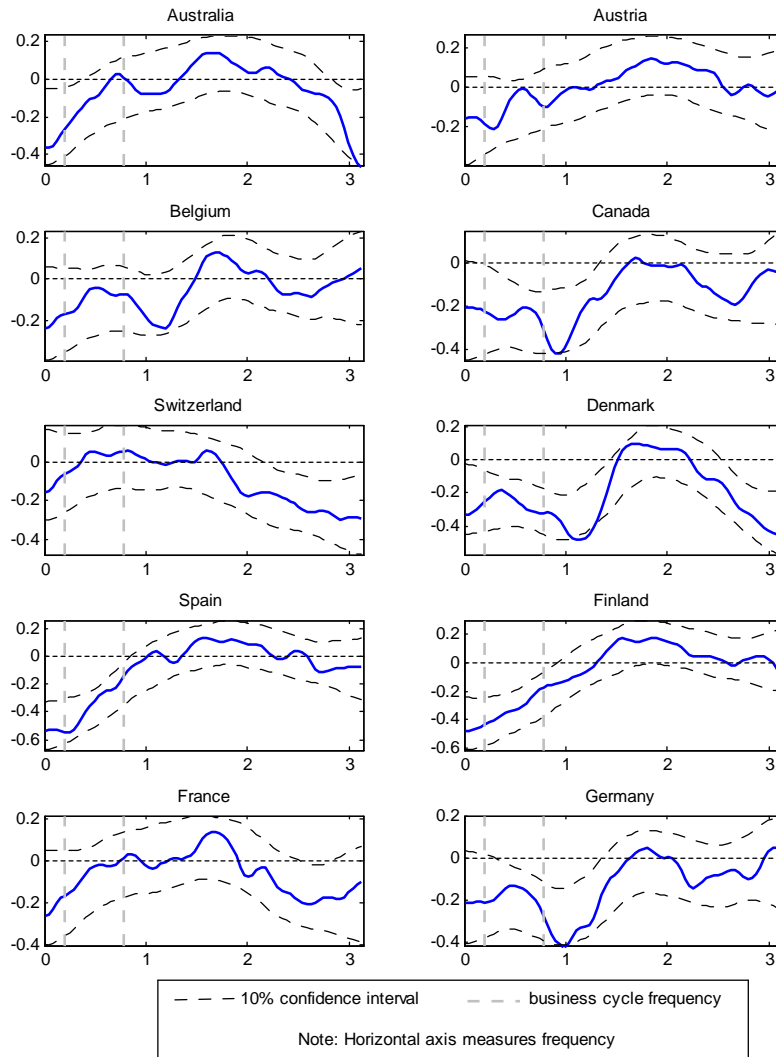
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D   Appendix. Figures and Tables

FIGURE 2A

Dynamic correlations of relative consumption and real exchange rates via-à-vis the US
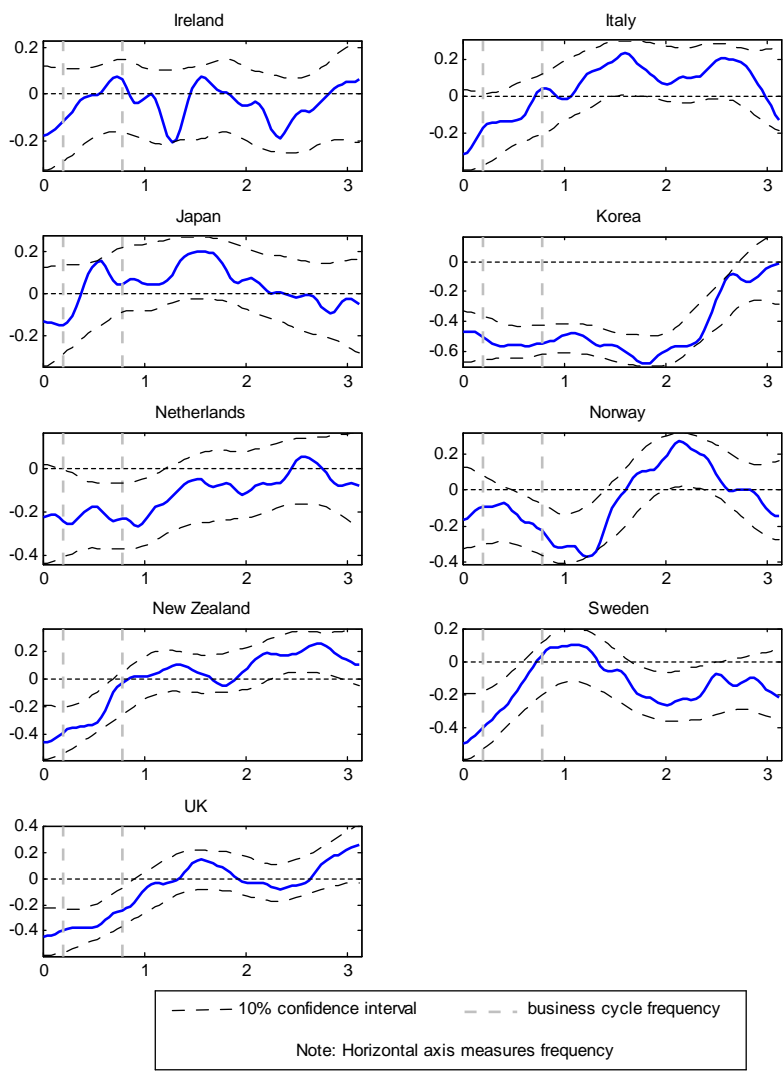
FIGURE 2B

Dynamic correlations of relative consumption and real exchange rates via-à-vis the US
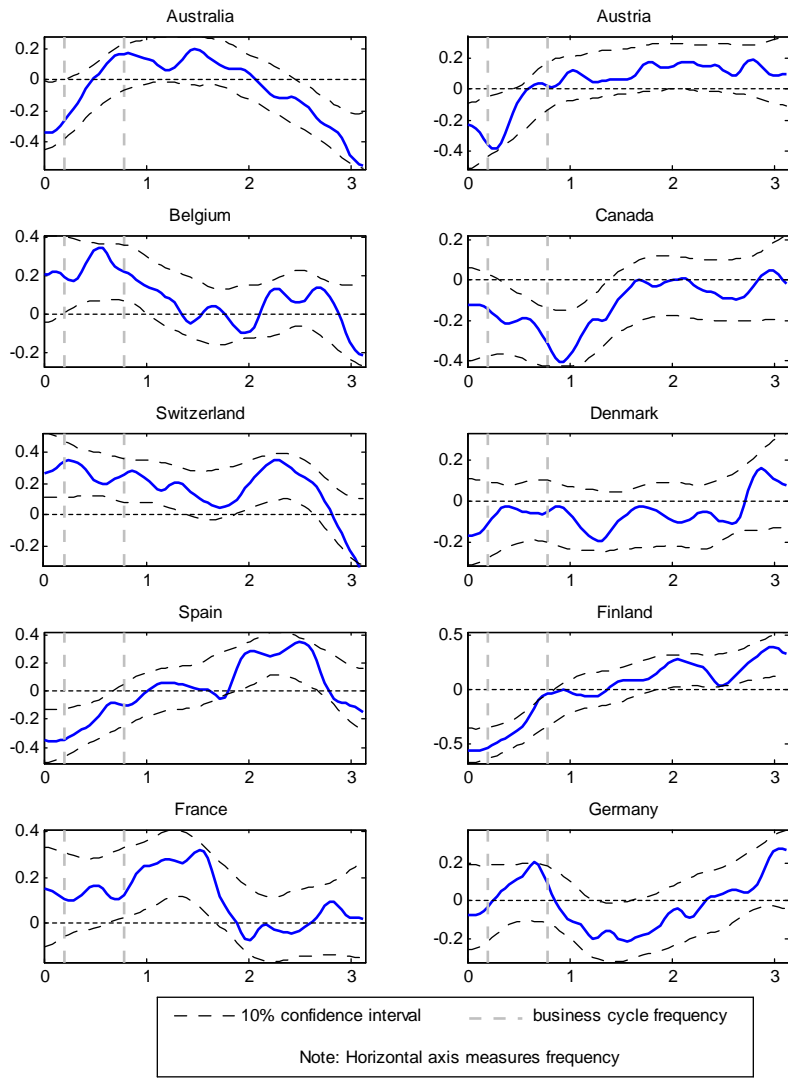
FIGURE 3A

Dynamic correlations of relative consumption and real exchange rates via-à-vis the rest-of-the-world
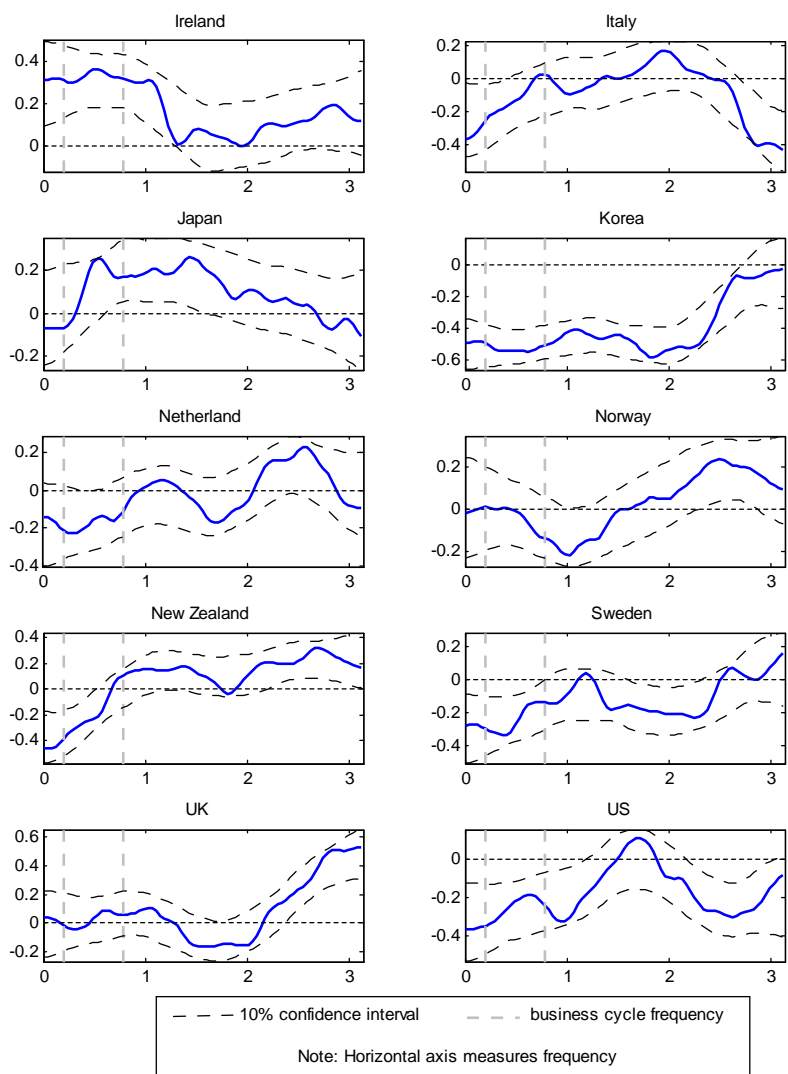
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FIGURE 3B

Dynamic correlations of relative consumption and real exchange rates via-à-vis the rest-of-the-world

TABLE 4

Dynamic correlations of differenced series at frequency zero
(cointegration test for variables in levels)

COUNTRY	Dynamic correlation (Rc,RER)	
	vs US	vs ROW
Australia	-0,36	-0,34
Austria	-0,16	-0,23
Belgium	-0,24	0,20
Canada	-0,20	-0,12
Switzerland	-0,16	0,27
Denmark	-0,33	-0,17
Spain	-0,53	-0,35
Finland	-0,48	-0,55
France	-0,26	0,15
Germany	-0,21	-0,07
Ireland	-0,18	0,31
Italy	-0,31	-0,37
Japan	-0,13	-0,07
Korea	-0,47	-0,49
Netherlands	-0,22	-0,14
Norway	-0,16	-0,02
New Zealand	-0,46	-0,46
Sweden	-0,49	-0,28
UK	-0,44	0,04
US	NA	-0,37

NOTE: Correlations equal to 1 or -1 signal cointegration of the series in levels.

E Appendix: the model economy

The world economy is specified after Corsetti et. al [2008]. It consists of two countries of equal size, H and F. Each country specializes in one type of tradable good, produced in a number of varieties or brands defined over a continuum of unit mass. Brands of tradable goods are indexed by $h \in [0,1]$ in the Home country and $f \in [0,1]$ in the Foreign country. In addition, each country produces an array of differentiated nontradables, indexed by $n \in [0,1]$. Nontraded goods are either consumed or used to make intermediate tradable goods h and f available to domestic consumers.

Firms producing tradable and nontraded goods are monopolistic suppliers of one brand of goods only. These firms combine capital with differentiated domestic labor inputs in a continuum of unit mass. Each worker occupies a point in this continuum, and acts as a monopolistic supplier of a differentiated type of labor input to all firms in the domestic economy. Firms operating in the distribution sector, by contrast, are assumed to operate under perfect competition. They buy tradable goods and distribute them to consumers using nontraded goods as the only input in production.

In the baseline model, prices will be assumed to be perfectly flexible. Alternative specifications include nominal price rigidities, via the assumption that firms face a quadratic cost of adjusting goods' prices. What follows describes the set up focusing on the Home country, with the understanding that expressions similar to the ones presented in the text also characterize the Foreign economy. As regards notation, variables referred to Foreign firms and households are marked with an asterisk.

E.1 The Household's Problem

The representative Home agent in the model maximizes the expected value of her lifetime utility, given by:

$$E \left\{ \sum_{t=0}^{\infty} U \left[C_t, \frac{M_{t+1}}{P_t}, L_t \right] \exp \left[\sum_{\tau=0}^{t-1} -\nu \left(C_\tau, \frac{M_{\tau+1}}{P_\tau}, L_\tau \right) \right] \right\}, \qquad (7)$$

where instantaneous utility U is a function of a consumption index, C_t, leisure, $(1 - L_t)$, and real money balances $\frac{M_{t+1}}{P_t}$. This recursive specification of preferences, according to which the discount factor is a function of past utility levels, guarantees the existence of a unique invariant distribution of wealth, independent of initial conditions.

Consumption baskets   Households consume all types of (domestically-produced) nontraded goods, and both types of traded goods. So $C_t(n)$ is consumption of brand n of Home nontraded good at time t; $C_t(h)$ and $C_t(f)$ are consumption of Home brand h and Foreign brand f. For each type of good, any individual brand is assumed to be an imperfect substitute for all other brands, with

constant elasticity of substitution θ_H and $\theta_N > 1$. Consumption of Home and Foreign goods by the Home household is defined as:

$$C_{H,t} \equiv \left[\int_0^1 C_t(h)^{\frac{\theta_H-1}{\theta_H}} dh \right]^{\frac{\theta_H}{\theta_H-1}}, \qquad C_{F,t} \equiv \left[\int_0^1 C_t(f)^{\frac{\theta_H-1}{\theta_H}} df \right]^{\frac{\theta_H}{\theta_H-1}}, \quad (8)$$

$$C_{N,t} \equiv \left[\int_0^1 C_t(n)^{\frac{\theta_N-1}{\theta_N}} dn \right]^{\frac{\theta_N}{\theta_N-1}}.$$

The full consumption basket, C_t, in each country is defined by the following CES aggregator

$$C_t \equiv \left[a_T^{1-\phi} C_{T,t}{}^\phi + a_N^{1-\phi} C_{N,t}{}^\phi \right]^{\frac{1}{\phi}}, \qquad \phi < 1, \tag{9}$$

where a_T and a_N are the weights on the consumption of traded and nontraded goods, respectively and $\dfrac{1}{1-\phi}$ is the constant elasticity of substitution between $C_{N,t}$ and $C_{T,t}$. The consumption index of traded goods $C_{T,t}$ is given by the following CES aggregator

$$C = C_T = \left[a_H^{1-\rho} C_H^\rho + a_F^{1-\rho} C_F^\rho \right]^{\frac{1}{\rho}}, \qquad \rho < 1. \tag{10}$$

Budget constraints and asset markets Home and Foreign agents trade an international bond, B_H, which pays in units of Home currency and is zero net supply. Households derive income from working, $W_t L_t$, from renting capital to firms, $R_t K_t$, from previously accumulated units of currency, and from the proceeds from holding the international bond, $(1+i_t)B_{H,t}$, where i_t is the nominal bond's yield, paid at the beginning of period t in domestic currency but known at time $t-1$. Households pay non-distortionary (lump-sum) net taxes T, denominated in Home currency, and use their disposable income to consume, invest in domestic capital, and buy bonds $B_{H,t+1}$. Only Home residents hold the Home currency, M_t. The individual flow budget constraint for the representative agent in the Home country is therefore:

$$M_{t+1} + B_{H,t+1} \leq M_t + (1+i_t)B_{H,t} + R_t K_t \tag{11}$$

$$+ \int_0^1 \Pi(h)dh + \int_0^1 \Pi(n)dn +$$

$$W_t L_t - T_t - P_{H,t}C_{H,t} - P_{F,t}C_{F,t} - P_{N,t}C_{N,t} - P_{INV,t}I_t$$

where $\int \Pi(h)dh + \int \Pi(n)dn$ is the share of profits from all firms h and n in the economy. The price indexes are as follows: $\overline{P}_{H,t}$ and $P_{H,t}$ denote the price of the Home traded good at the *producer* and *consumer* level, respectively, $P_{F,t}$ is the consumer price of Home imports; $P_{N,t}$ is the price of nontraded goods; P_t is the consumer price index; $P_{INV,t}$ is the price of investment. As the international bonds are assumed to be in zero net supply, market clearing implies $B_{H,t+1} = -B_{H,t+1}^*$.

Investment is a Cobb-Douglas composite of tradable and nontradable goods, in line with the evidence in Bems [2005]. The tradable component of investment combines local and imported goods based on the same CES aggregator as tradable consumption. Different from tradable consumption, however, investment is not subject to distribution services. Observe that this specification of investment corresponds to the notion that intermediate imported inputs contribute to the formation of capital in the economy. The law of motion for the aggregate capital stock is given by:

$$K_{t+1} = I_t + (1-\vartheta)K_t + \frac{b}{2}\left(\frac{I_t}{K_t} - \vartheta\right)^2, \tag{12}$$

where b is an adjustment cost parameter, as in CKM. The capital stock, K, can be freely reallocated between the traded (K_H) and nontraded (K_N) sectors. The depreciation rate is denoted by ϑ.

The household's problem then consists of maximizing lifetime utility, defined by (7), subject to the constraints (11) and (12).

E.2 Firms' problem

Firms producing Home tradables (H) and Home nontradables (N) are monopolist in their variety of good; they employ a technology that combines domestic labor and capital inputs, according to the following Cobb-Douglas functions:

$$
\begin{aligned}
Y(h) &= Z_H K(h)^{1-\xi} L(h)^{\xi} \\
Y(n) &= Z_N K(n)^{1-\zeta} L(n)^{\zeta},
\end{aligned}
\tag{13}
$$

where Z_H and Z_N are sectoral random disturbance following a statistical process to be determined below. Capital and labor are freely mobile across sectors.

The specification of the distribution sector is in the spirit of the factual remark by Tirole ([1995], page 175) that "production and retailing are complements, and consumers often consume them in fixed proportions". As in Burstein, Neves and Rebelo [2003] and Corsetti and Dedola [2005], bringing one unit of traded goods to consumers requires η units of a basket of differentiated nontraded goods:

$$\eta = \left[\int_0^1 \eta(n)^{\frac{\theta_N-1}{\theta_N}}dn\right]^{\frac{\theta_N}{\theta_N-1}}. \tag{14}$$

Flexible prices With flexible prices, firms hire labor and capital from households to maximize their profits:

$$
\begin{aligned}
\pi_t(h) &= \overline{p_t}(h)D_t(h) - W_t L_t(h) - R_t K_t(h) \\
\pi_t(n) &= p_t(n)D_t(n) - W_t L_t(n) - R_t K_t(n)
\end{aligned}
\tag{15}
$$

where $\bar{p}_t(h)$ is the *wholesale* price of the Home traded good and $p_t(n)$ is the price of the nontraded good. W_t denote the aggregate wage rate, while R_t represents the capital rental rate.

Consider first the optimal pricing problem faced by firms producing nontradables for the Home market. The demand for their product is

$$D(n) + \eta(n) = [p_t(n)]^{-\theta_N} P_{N,t}^{\theta_N} \left[D_{N,t} + \eta \left(\int_0^1 D_t(h)dh + \int_0^1 D_t(f)df \right) \right], \quad (16)$$

where $D_{N,t}$ is the (consumption and investment) aggregate demand for nontraded goods. It is easy to see that their optimal price will result from charging a constant markup over marginal costs:

$$
\begin{aligned}
p_t(n) &= P_{N,t} = \frac{\theta_N}{\theta_N - 1} MC_{N,t} \\
&= P_{N,t} = \frac{\theta_N}{\theta_N - 1} \frac{W_t^\zeta R_t^{1-\zeta}}{Z_{N,t}}
\end{aligned}
\quad (17)
$$

Now, let $\bar{p}_t(h)$ denote the price of brand h expressed in the Home currency, at *producer* level. With a competitive distribution sector, the consumer price of good h is simply

$$p_t(h) = \bar{p}_t(h) + \eta P_{N,t}. \quad (18)$$

In the case of firms producing tradables, "pricing to market" derives endogenously from the solution to the problem of the Home representative firm in the sector:

$$Max_{\bar{p}(h),\bar{p}^*(h)} \quad [\bar{p}_t(h)D_t(h) + \mathcal{E}_t\bar{p}_t^*(h)D_t^*(h)] - \frac{W_t^\xi R_t^{1-\xi}}{Z_{H,t}} [D_t(h) + D_t^*(h)] \quad (19)$$

where $\mathcal{E}_t$ is the nominal exchange rate, expressed in home currency units, and

$$D_t(h) = \left(\frac{P_{H,t}}{\bar{p}_t(h) + \eta P_{N,t}} \right)^{\theta_H} C_{H,t}, \quad D_t^*(h) = \left(\frac{P_{H,t}^*}{\bar{p}_t^*(h) + \eta P_{N,t}^*} \right)^{\theta_H^*} C_{H,t}^*. \quad (20)$$

Making use of (17), the optimal wholesale prices for the consumption good $\bar{p}(h)$ and $\bar{p}^*(h)$ are:

$$\bar{p}_t(h) = \frac{\theta_H}{\theta_H - 1} \left(1 + \frac{\eta}{\theta_H} \frac{\theta_N}{\theta_N - 1} \frac{Z_{H,t}}{Z_{N,t}} \frac{W_t^\zeta R_t^{1-\zeta}}{W_t^\xi R_t^{1-\xi}} \right) \frac{W_t^\xi R_t^{1-\xi}}{Z_{H,t}} = mk_{H,t} \frac{W_t^\xi R_t^{1-\xi}}{Z_{H,t}},$$
$$(21)$$
$$\mathcal{E}_t\bar{p}^*(h) = \frac{\theta_H^*}{\theta_H^* - 1} \left(1 + \frac{\eta}{\theta_H^*} \frac{\theta_N^*}{\theta_N^* - 1} \frac{Z_{H,t}}{Z_{N,t}^*} \frac{\mathcal{E}_t W_t^{*\zeta} R_t^{*1-\zeta}}{W_t^\xi R_t^{1-\xi}} \right) \frac{W_t^\xi R_t^{1-\xi}}{Z_{H,t}} = mk_{H^*,t} \frac{W_t^\xi R_t^{1-\xi}}{Z_{H,t}},$$
$$(22)$$

where $mk_{H,t}$ and $mk_{H^*,t}$ denote the markups. Unlike the case of nontraded goods (17), in this case the markups charged by the Home firms include a state-contingent component — in brackets in the above expression — that varies as a function of productivity shocks, monetary innovations (affecting the exchange rate) and relative wages. Since in general $mk_{H,t}$ will not equal $mk_{H^*,t}$, even

when $\theta_H^* = \theta_H$, the optimal wholesale price of tradable goods will not obey the law of one price $(\bar{p}_t(h) \neq \mathcal{E}_t \bar{p}_t^*(h))$. This result reflects the difference in the elasticity of demand faced by the upstream monopolist at Home and abroad brought about by any asymmetry in relative productivity and/or factor prices.

Finally, note that since there are no distribution costs in investment, the price of investment goods in a flexible-price allocation will be equal to the standard expression without the state contingent component of markups.

Sticky Prices Following Rotemberg [1982] and Dedola and Leduc [2001], firms in the traded and non-traded goods sectors face a quadratic cost when adjusting their prices (a cost which is however set equal to zero in steady state). To change its product prices, a firm needs to use resources, in the form of a CES basket of all the goods in the same sector of the economy; thus the price-adjustment costs faced by firms in the traded and non-traded goods sector are respectively:

$$AC_{H,t}^p (h) = \frac{\kappa_H^p}{2} \left(\frac{\bar{p}_t(h)}{\bar{p}_{t-1}(h)} - \pi \right)^2 D_{H,t}, \qquad AC_{H,t}^{p*} (h) = \frac{\kappa_H^p}{2} \left(\frac{\bar{p}_t^*(h)}{\bar{p}_{t-1}^*(h)} - \pi^* \right)^2 D_{H,t}^*,$$

$$(23)$$

and

$$AC_t^p (n) = \frac{\kappa_N^p}{2} \left(\frac{p_t(n)}{p_{t-1}(n)} - \pi \right)^2 D_{N,t}, \tag{24}$$

where π and π^* are the domestic and foreign steady-state inflation rates and κ_H^p, κ_H^p, and κ_N^p are adjustment cost parameters. As firms producing traded goods charge different prices in different markets, the cost of changing prices is incurred in each market independently of the other. Note that, innocuously, we assume that both $AC_{H,t}^p (h)$ and $AC_{H,t}^{p*} (h)$ are denominated in units of domestic traded goods.

E.3   Price indexes and monetary policy

Distribution costs create a wedge between the producer price and the consumer price of each good. Since firms in the distribution sector are perfectly competitive, the consumer price of the Home traded good $P_{H,t}$ is simply the sum of the price of Home traded goods at producer level $\overline{P}_{H,t}$ and the value of the nontraded goods that are necessary to distribute it to consumers

$$P_{H,t} = \overline{P}_{H,t} + \eta P_{N,t}. \tag{25}$$

Hereafter the price index of tradables and the utility-based CPIs will be denoted as:

$$P_{T,t} = \left[a_H P_{H,t}^{\frac{\rho}{\rho-1}} + a_F P_{F,t}^{\frac{\rho}{\rho-1}} \right]^{\frac{\rho-1}{\rho}} \tag{26}$$

$$P_t = \left[a_T P_{T,t}^{\frac{\phi}{\phi-1}} + a_N P_{N,t}^{\frac{\phi}{\phi-1}} \right]^{\frac{\phi-1}{\phi}}.$$

Foreign prices, denoted with an asterisk and expressed in the same currency as Home prices, are similarly defined. In both countries, inflation π_t (π_t^*) is conventionally defined as the rate of growth of the CPIs, P and P^*.

The presentation of the model is then completed with the specification of monetary policy, which pins down nominal variables. It is assumed that monetary authorities follows a Taylor-type rule, taking the following form:

$$i_t = \rho i_{t-1} + \Psi(1-\rho)E(\pi_{t+1} - \pi^{ss}) + \gamma(1-\rho)(y_t - y^{ss}). \qquad (27)$$

Monetary authorities set the short-term nominal interest rate, i_t, as a function of the deviations of expected CPI inflation (π) and GDP (y) from steady state values (π^{ss} and y^{ss}).

E.4 Equilibrium and exchange rates

A competitive equilibrium for the world economy presented above is defined along the usual lines, as a set of processes for quantities and prices in the Home and Foreign country satisfying: (i) the household and firms optimality conditions; (ii) the market clearing conditions for each good and asset, including money; (iii) the appropriate resource constraints — whose specification can be easily derived from the above and is omitted to save space.

Before delving into the analysis, it is useful to provide details on how the exchange rate is determined in nominal and real terms — the real exchange rate being customarily defined as the ratio of consumption prices across countries, expressed in the same currency, $\mathcal{E}P^*/P$. The crucial equilibrium condition for real exchange rate determination is obtained by combining the standard Euler equations for bond holdings by the Home and the Foreign households:

$$R_t^{-1} = \widetilde{\beta}_t E_t \left[\frac{\partial U\left[C_{t+1}, \frac{M_{t+2}}{P_{t+1}}, L_{t+1}\right]/\partial C_{t+1}}{\partial U\left[C_t, \frac{M_{t+1}}{P_t}, L_t\right]/\partial C_t} \frac{P_t}{P_{t+1}} \right] = \widetilde{\beta}_t^* E_t \left[\frac{\partial U\left[C_{t+1}^*, \frac{M_{t+2}^*}{P_{t+1}^*}, L_{t+1}^*\right]/\partial C_{t+1}^*}{\partial U\left[C_t^*, \frac{M_{t+1}^*}{P_t^*}, L_t^*\right]/\partial C_t^*} \frac{P_t^*}{P_{t+1}^*} \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}} \right]$$

$$(28)$$

where $\widetilde{\beta}_t$ ($\widetilde{\beta}_t^*$) denote the endogenous discount factor implicitly defined in (7). This standard condition is the same as the one discussed in the text.

With flexible prices, given the money demand function implicit in (7), monetary policy pins down the evolution of the price level and the other nominal variables in each country; thus, given the equilibrium real exchange rate, the nominal exchange rate will be determined by the relative monetary stance in the countries. With sticky prices, instead, monetary policy will have also some short-run effects on real variables; however, in line with the Taylor rule (27), monetary policy will be mainly concerned with stabilizing inflation, so that price level movements are quite smooth. It follows that nominal exchange rates will closely mimic real exchange rates.

E.5 Calibration

The parameters invariant across the three calibrations used in the text are listed in Table 5. The parameters of the productivity process are instead as follows. Based on the OECD STAN database, the autocorrelation matrix is

$$\lambda = \begin{vmatrix} 0.82 & -0.06 & 0.10 & 0.24 \\ -0.06 & 0.82 & 0.24 & 0.10 \\ -0.02 & 0.02 & 0.96 & 0.01 \\ 0.02 & -0.02 & 0.01 & 0.96 \end{vmatrix}$$

and the variance-covariance matrix is

$$\Omega = \begin{vmatrix} 0.047 & 0.022 & 0.009 & 0.004 \\ 0.022 & 0.047 & 0.004 & 0.009 \\ 0.009 & 0.004 & 0.009 & -0.001 \\ 0.004 & 0.009 & -0.001 & 0.009 \end{vmatrix}.$$

In CDL-low elasticity, the elasticity of substitution between domestic and foreign tradables is set equal to 0.74. Because of distribution costs, the trade elasticity is one half the elasticity of substitution.

In our second specification (CDL, high elasticity), the elasticity of substitution between domestic and foreign tradables is set equal to 8. The shocks to the tradable good sector now have an autoregressive parameter of 0.99 — while spillovers are set to zero to guarantee stationarity.

For the specification reproducing Benigno and Thoenissen (2006), the auto-correlation matrix is set as follows

$$\lambda = \begin{vmatrix} 0.95 & 0 & 0.22 & 0 \\ 0 & 0.95 & 0 & 0.22 \\ 0 & 0 & 0.30 & 0 \\ 0 & 0 & 0 & 0.30 \end{vmatrix}$$

while the variance-covariance matrix is

$$\Omega = \begin{vmatrix} 0.037 & 0.015 & 0.072 & 0.044 \\ 0.015 & 0.037 & 0.044 & 0.072 \\ 0.072 & 0.044 & 0.0025 & 0.021 \\ 0.044 & 0.072 & 0.021 & 0.0025 \end{vmatrix}.$$

As explained in the text, we have changed these processes relative to the BT contribution, as to enable our model to match the transmission mechanism and a negative BS correlation as envisioned by BT.

Table 5. Parameter values

Benchmark Models	
Preferences and Technology	
Risk aversion	$\sigma = 2$
Disutility of labor	$\alpha = 2.03,\ 4.11$
Velocity parameter	$\chi = 0.1$
Elasticity of substitution between:	
Home and Foreign traded goods	$\frac{1}{1-\rho} = 7.4, 1.5, 2, 8$
traded and non-traded goods	$\frac{1}{1-\phi} = 0.74$
Home non-traded goods	$\theta_{\mathrm{N}} = 7.7$
Home traded goods	$\theta_{\mathrm{H}} = 15.3$
Elasticity of the discount factor	
with respect to C and L	$\psi = 0.005,\ 0.003$
Distribution margin	$\mu = 0.5$
Labor share in tradables	$\xi = 0.61, 0.33$
Labor share in nontradables	$\zeta = 0.56$
Depreciation rate	$\vartheta = 0.025$
Monetary Policy	
Lagged interest-rate coefficient	$\rho = 0.84$
Weight on inflation	$\Psi = 2.19$
Weight on output gap	$\gamma = 0.3$