Borders and Nominal Exchange Rates in Risk-Sharing*

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Abstract

Models of risk-sharing predict that relative consumption growth rates are positively related to changes in real exchange rates. We investigate this hypothesis using a new multi-country and multi-regional data set. Within countries, we find evidence for risk-sharing: episodes of high relative regional consumption growth are associated with regional real exchange rate depreciation. Across countries however, the association is reversed: relative consumption and real exchange rates are negatively correlated. We define this reversal as a ‘border’ effect. We find the border effect and show that it accounts for over half of the deviations from full risk-sharing. Since cross-border real exchange rates involve different currencies, it is natural to ask how much of the border effect is accounted for by movements in exchange rates. Our measures indicate that a large part the border effect comes from nominal exchange rate fluctuations. We develop a simple open economy model that is consistent with the importance of nominal exchange rate fluctuations.

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variability in accounting for deviations from cross-country risk-sharing.

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## 1 Introduction

Many studies have documented the failure of naive models of consumption risk-sharing. This is true both within countries (risk-sharing across provinces or states) and across countries. Recognizing that consumption prices differ across time and locations leads to a more elaborate test for consumption risk-sharing, incorporating both within and between-country real exchange rate movements. This extended model predicts that relative consumption growth rates (across regions or countries) are highly positively correlated with movements in real exchange rates. In international data, it is well known that this prediction is strongly refuted. This is essentially the well-known ‘Backus-Smith’ puzzle ([Backus and Smith](1993)). The evidence in fact seems to indicate that relative consumption and real exchange rates are negatively correlated across countries. Numerous recent studies have attempted to rationalize the source of this failure of the basic model of international risk-sharing.

To assess the failure of full international risk-sharing in the data it is critical to develop a benchmark for comparison. The degree of within-country risk-sharing provides a simple and model-free benchmark. In this paper we revisit the relationship between relative consumption and real exchange rates by focusing on the contrast between risk-sharing within countries and between countries. Although previous studies have noted the substantial differences in measures of risk-sharing within and across countries (see [Hess and Van Wincoop](2000) for an overview), there has been little investigation of the predictions of risk-sharing for the consumption-real exchange

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1 See, for instance, [Obstfeld](2007), [Chari, Kehoe, and McGrattan](2002), [Corsetti, Dedola, and Leduc](2008), [Benigno and Thoenissen](2008), [Kollmann](2012) and others.
rate relationship within countries. Our paper, to our knowledge, provides the first comprehensive evidence on this within-country relationship.

We use a new multi-country and multi-regional data set on consumption and prices indices. We find a sharp dichotomy between within-country and across-country comparisons. Across countries, as in previous studies, we find that relative consumption growth and real exchange rates are negatively correlated. Within-country, however, relative consumption growth and real exchange rates tend to be positively correlated, as implied by standard models of risk-sharing. Because our measures links a large number of bilateral locations within and between countries, we can identify a ‘border effect’ in the consumption-real exchange rate relationship, similar to the border effect Engel and Rogers (1996). Our first main result is that the border effect reverses the sign of the consumption-real exchange rate relationship.

What accounts for the border effect? Cross-border real exchange rates involve different currencies, so an obvious candidate explanation is that cross-country real exchange rates behave in a different way than within-country real exchange rates due to fluctuations in the nominal exchange rate. Following this logic, and using observed nominal exchange rates, we can decompose the border effect into that accounted for by the nominal exchange rate, and a residual, orthogonal to exchange rate movements. We find that movements in the nominal exchange rate account for a large part of the consumption-real exchange rate relationship.

Since the negative cross-border correlation of consumption and real exchange rates is to a large extent accounted for by the nominal exchange rate, we then ask whether eliminating nominal exchange rate variability improves risk-sharing outcomes. Such a decomposition is tricky as cross-

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2 An extensive literature in open economy macroeconomics, going back at least to Mussa (1986), has emphasized the large and puzzling behaviour of nominal exchange rates in accounting for real exchange rate variability. Recent papers by Hadzi-Vaskov (2008) and Hess and Shin (2010) relate the anomalous findings on the consumption-real exchange rate relationship to the movements of the nominal exchange rate. They show that in OECD countries, the negative correlation between relative consumption and real exchange rates is proximately accounted for by movements in nominal exchange rates. We discuss these papers more extensively below.
border comparisons are almost always associated with some nominal exchange rate movements. A negative consumption-real exchange rate correlation may be an implication of national borders, but in fact independent of exchange rate policy. One way to investigate this is to take cross-border pairs which exhibit little exchange rate variability, and compare their properties with those cross-border pairs whose exchange rate is freely floating. Our sample includes both Germany and Spain, which since 1999 have had no bilateral nominal exchange rate variability, and in fact even before this, to the beginning of the sample (1995), had very small bilateral nominal exchange rate volatility. We find that the consumption-real exchange rate correlation for Germany-Spain bilateral pairs is positive, while it remains negative and significant between all other bilateral pairs. We then further extend this exercise to a set of 12 Eurozone countries and contrast the consumption-real exchange rate correlation among them during the pre-euro period and euro-period. We show that the average bilateral correlation is overwhelmingly negative for all countries, except Greece, in the pre-Euro period of 1974-1998, equal to -0.19 across all 12 countries. Importantly, the correlation turns positive for the majority of countries after the introduction of the Euro. In two out of four countries where the correlation remained negative (France and Italy), its value significantly increased relative to the pre-Euro period. Overall, the average post-Euro correlation across countries is positive 0.05. This offers some evidence for the importance of the nominal exchange rate regime itself, as opposed to indirect effects of national borders, in determining the consumption-real exchange rate correlation.

To assess the contribution of the border to the failure of consumption risk-sharing we compute a cross-section of deviations from full risk-sharing between all bilateral locations, over the whole sample. Our previous results suggest that this should be positively correlated with the presence of a national border. Indeed we find that the border accounts for 53% of the deviation from full

\footnote{This is reminiscent of the Stockman (1987) argument that real exchange rate volatility may be determined by changes in real fundamentals which are only realized following changes in exchange rate regimes.}
risk-sharing across regions. To evaluate the contribution of the nominal exchange rate to the border effect we compute a proxy real exchange rate assuming that the nominal exchange rate is equal to the sample average for all bilateral pairs. We find that in this case the border effect on the risk-sharing measure declines by one-third, suggesting a non-trivial role for the nominal exchange rate in explaining the deviations from full risk-sharing.

Our results present a challenge to the existing models of real exchange rate determination. A substantial literature argues that the presence of incomplete financial markets may explain why relative consumption growth and the real exchange rate co-vary negatively in the data. See, for instance, Corsetti, Dedola, and Leduc (2008) and Benigno and Thoenissen (2008). However, for the most part, this literature offers no explicit role for the nominal exchange rate regime in resolving the puzzle. But, if nominal goods prices are sticky, the correlation of relative consumption and the real exchange rate may depend on the degree of nominal exchange rate flexibility. However, many sticky price models (e.g. Clarida, Gali, and Gertler 2002; Devereux and Engel 2003) exhibit volatile real and nominal exchange rates, but still have the implication that the cross country risk sharing condition between consumption and real exchange rates holds exactly if financial markets are complete.

Our model combines features of the previous literature on the Backus-Smith puzzle with a two country ‘New Keynesian’ model with gradual price adjustment. We also introduce heterogeneous regions with two sectors (tradable and nontradable) within each country to study intra-national risk-sharing and its determinants. Across countries the model admits flexible or fixed exchange rate regime. We then posit two key questions to the model: (i) whether this model exhibits the

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4 Most proposed resolutions emphasize the joint role of incomplete markets and shocks which generate strong income effects. The intuition is that a country which has a faster growing consumption experiences an appreciating real exchange rate. But when we extend the anomaly to encompass both regions within a country, and the importance of the nominal exchange rate, these explanations are not complete, since in these models, the nominal exchange rate has no implication for the consumption-real exchange rate correlation. These models would predict that consumption-real exchange rate correlation is the same across countries and across regions within countries.
property that intra-national risk-sharing is higher than international risk-sharing; (ii) whether the sign of the consumption-real exchange rate correlation across countries depends upon the exchange rate regime.

We find that when the model is calibrated using standard parameters, and there exists a degree of ‘rule of thumb pricing’, as in Gali and Gertler (1999), the empirical findings above can be rationalized. Within countries, the model predicts a consistently positive consumption-real exchange rate correlation, but it is far from the complete markets implication of a unit correlation. However, this correlation rises in the presence of endogenous fiscal transfers across regions within a country and higher co-movement between productivity shocks within countries. For the international correlation, the key insight that the model brings is that when the exchange rate is flexible, traded goods productivity shocks can lead both the nominal and real exchange rate and relative consumption to move in different directions. Nontraded productivity shocks produce a positive correlation between relative consumption and the real exchange rate. Since the traded productivity shocks are larger in the data and in our calibration, the unconditional cross-country consumption-real exchange rate correlation is negative under the flexible exchange rate regime. When the exchange rate is fixed, the impact on the real exchange rate is significantly dampened, and more so following traded productivity shocks. In our baseline model, the difference is enough to reverse the sign of the cross-country consumption-real exchange rate correlation when nominal exchange rates are held fixed.

Rule of thumb behavior in price setting is critical however. The reason is that the standard model, specified according to the Calvo price setting mechanism with full forward looking pricing, allows for too much immediate price response. Even consequent upon fixing the exchange rate, in the Calvo model, prices respond sufficiently flexibly that the real exchange rate-relative consumption correlation is little changed when comparing floating and fixed exchange rates. With rule of thumb
pricing, as in Gali and Gertler (1999), prices show much less initial response to shocks, and in the calibrated benchmark model, this leads to a substantial difference in the consumption-real exchange rate correlation between fixed and floating exchange rates.\footnote{This result is very similar to the well known critique of the Calvo model that it allows too much variability in inflation in response to new information - see Gali and Gertler (1999), Gali, Gertler, and Lopez-Salido (2001), McAdam and Willman (2004), and Christiano, Eichenbaum, and Evans (2005). Alternative approaches to introducing sluggishness in inflation response, such as Mankiw and Reis (2002), or Christiano, Eichenbaum, and Evans (2005), who assume that prices must be set before current information on shocks is released, would produce very similar results to those we find below.}

Our analysis is related to a recent empirical literature on international risk-sharing and the real exchange rate. The most closely related papers are those by Hadzi-Vaskov (2008) and Hess and Shin (2010) who investigate the sources of the Backus-Smith puzzle in cross-country data. Hadzi-Vaskov (2008) shows that movements in the nominal exchange rate contribute significantly to the negative consumption-real exchange rate correlation in a sample of Eurozone countries. Hess and Shin (2010) reach the same conclusion in a sample of OECD countries. Our paper builds on theirs, but differs in several respects. In particular, by using intra-national data for numerous countries, we can define a ‘border effect’ in the relationship between consumption and real exchange rates.\footnote{Hess and Shin (2010) estimate the consumption real exchange rate correlation in intra national data for the US states only. Therefore, they can not contrast their findings directly with the international evidence to isolate the role of the border.}

In addition, in our analysis we employ several decomposition approaches that allow us to robustly assess and quantify the role played by the border and the nominal exchange rate in the deviation from risk-sharing. Finally, we provide an explicit theoretical analysis of our results.

The rest of the paper is organized as follows. The next section presents the empirical evidence on the role of the nominal exchange rate in the consumption-real exchange rate relationship. In Section 3 we develop models with sticky prices to interpret our empirical findings. Section 4 concludes.
2 Estimating the border effect

2.1 Key theoretical relationship

Consider a multi-jurisdiction (where a jurisdiction may be a country or region) stochastic model. The utility of a representative household in jurisdiction \( j = 1, \ldots, J \) is given by:

\[
E_t \sum_{s=0}^{\infty} \beta^s U(C_{j,t+s}), \quad \beta < 1
\]

where \( \beta \) is the subjective discount factor, \( C_{j,t} \) denotes a composite consumption good in country \( j \). \( E_t \) is period-\( t \) expectation. Define \( P_{j,t} \) to be the price of a representative consumption basket in jurisdiction \( j \) in period \( t \). Also let \( S_{i,j}^t \) be the exchange rate that converts prices from country \( j \)’s currency to country \( i \)’s currency in period \( t \). If jurisdictions are within the same country, then \( S_{i,j}^t = 1 \). Then the real exchange rate between any two regions \( i \) and \( j \) in different countries is given by \( RER_{t}^{ij} = S_{i,j}^t P_{j,t} / P_{i,t} \), or \( RER_{t}^{ij} = P_{j,t} / P_{i,t} \) if \( i \) and \( j \) are two regions in the same country.

Suppose that there is a complete set of state-contingent securities available to households in all countries. In this case, the key optimality condition is to equate marginal utilities of consumption across countries (or regions), adjusted for differences in price levels, evaluated in a common currency:

\[
U_c(C_{i,t})RER_{t}^{ij} = U_c(C_{j,t}). \tag{2.1}
\]

This equation must hold in every date and state of the world, between any two countries or regions \( i \) and \( j \). It says that in equilibrium, consumption between households \( i \) and \( j \) must be allocated in a way that marginal utility (converted into the same units using the real exchange rate) is equalized.

\footnote{It is possible that the consumption real exchange rate relationship is affected by country or region-specific preference shocks which directly alter the marginal utility of consumption, apart from consumption itself. As in most of the literature, we abstract from such shocks, since they are unobservable, and allow for a rationalization of a consumption-real exchange rate correlation of any sign or size.}
across countries. Then if utility is of a constant relative risk aversion (CRRA) form, with the
coefficient of relative risk aversion \( \sigma \), equation (2.1) becomes

\[
\left( \frac{C_{i,t}}{C_{j,t}} \right)^\sigma = RER_{t}^{j,i},
\]

or equivalently in logs

\[
\sigma (\ln C_{i,t} - \ln C_{j,t}) = \ln RER_{t}^{j,i}.
\]

Following Backus and Smith (1993) we write the expression above in growth rates:

\[
\sigma (\Delta \ln C_{i,t} - \Delta \ln C_{j,t}) = \Delta \ln RER_{t}^{j,i}, \tag{2.2}
\]

where \( \Delta \ln X_{i,t} = \ln X_{i,t} - \ln X_{i,t-1} \). We should note that taking first-differences provides a natural
de-trending of the consumption and real exchange rate variables. We also use the Hodrick-Prescott
(HP) filter to de-trend the variables, as in Corsetti, Dedola, and Leduc (2008). This provides us
with a robustness check as it allows us to focus on variations at business cycle frequency. We find
that all results remain robust to this alternative de-trending, and therefore in what follows we focus
on the results in growth rates as they correspond the closest to the approach in Backus and Smith
(1993).\footnote{The estimation results with HP-filtered data are provided in the online appendix.}

These expressions establish the close relationship between the real exchange rate and relative
consumption between jurisdictions \( i \) and \( j \). In particular, equation (2.2) implies that consumption
growth between \( t - 1 \) and \( t \) should be relatively higher in jurisdictions whose real exchange rates
depreciate during the same period. Therefore, if markets are complete, the correlation, \( \rho_{e,c_{j}/c_{i}} =
\text{corr}(\Delta \ln RER_{t}^{i,j}, \sigma \Delta \ln \frac{C_{i,t}}{C_{j,t}}) \), should be equal to 1, as pointed out by Backus and Smith (1993)
and Kollmann (1995). A version of condition (2.2), defined in terms of conditional expectations,
will also hold even under incomplete markets, so long as some financial assets can be traded across
countries (see e.g. Obstfeld and Rogoff (1996)). Notice that if relative purchasing power parity
(PPP) holds, so that real exchange rate is constant, then $\Delta \ln RER_{t}^{j,i} = 0$. In this case we get a
standard risk-sharing result that consumption growth rates should be equal across jurisdictions.
This simple implication has been tested extensively in the cross-country context in Asdrubali,
Sørensen, and Yosha (1996), Athanasoulis and van Wincoop (2001), Bayoumi and Klein (1997),

2.2 Empirical facts

Equation (2.2) gives us the key testable relationship implied by the model. As is clear from (2.2),
the condition can be applied to any two locations of interest: countries, regions, states/provinces/
prefectures, etc. We use this relationship to study the extent of national and regional risk-sharing
between the US, Canada, Germany, Japan and Spain.

We employ intra-national data on consumption, output and prices in 50 US states during 1969-2006, 12 Canadian provinces and territories
during 1981-2007, 47 Japanese prefectures over 1990-2005 period, 16 German bundeslaender over
Using this
data we compute all possible unique bilateral pairs of differences between consumption, price and
output growth rates. The summary statistics for our dataset are reported in Appendix A.2.

To simplify notation we will use $\Delta c_{t}^{i,j}$ to denote relative consumption growth between two
locations $i$ and $j$, so that $\Delta c_{t}^{i,j} = \Delta \ln C_{i,t} - \Delta \ln C_{j,t}$; and $\Delta e_{t}^{j,i}$ to denote real exchange rate growth
between locations $i$ and $j$, so that $\Delta e_{t}^{j,i} = \Delta \ln RER_{t}^{j,i}$. Then based on equation (2.2) we posit the

9In the online appendix we also study the US and Canada separately since these two countries are often the focus of
the studies of risk-sharing and the law of one price deviations (see Engel and Rogers (1996), Gorodnichenko and
Tesar (2009)).

10 These are the countries for which historical jurisdiction-level data on consumption and prices exist. For the US
we use retail sales to proxy for private consumption. Consumption and GDP are real and in per capita terms. Details
on data sources and series construction are provided in the Appendix A.1.
following specification to link relative consumption growth and real exchange rate growth:

$$\Delta c_{i,j}^t = \beta_0 + \beta_1 \Delta c_{i}^t + \beta_2 (\Delta e_{i}^t \ast \text{border}_{i,j}) + v_{i,j}^t,$$

where $v_{i,j}^t$ is the error term arising due to preference shocks, measurement error, etc. $\text{border}_{i,j}$ is the border dummy that takes value of one for all cross-border location pairs, and zero otherwise. This allows us to distinguish between cross-country risk sharing and within-country risk sharing.

This specification restricts the relationship between the real exchange rate and relative consumption to be the same for any two locations within countries or any two locations across countries. However, it is plausible to posit that the same change in the real exchange rate could be associated with different movements in relative consumption depending on the particular locations observed. For instance, there may be differences in the degree of openness in goods or financial markets between two jurisdictions that are not reflected in changes in the real exchange rate. Distance represents a natural explanatory variable in the studies of the deviations from the law of one price between location pairs. In terms of deviations from risk sharing, distance may seem somewhat less compelling, since a) it may already be incorporated in the movement in real exchange rates, and b) it is a constant, and may not affect the risk sharing relationship when measured in growth rates. Nevertheless, some studies (e.g. Portes and Rey (2005), Okawa and van Wincoop (2010)) have documented the explanatory power of gravity type variables in accounting for international financial activity. To allow for this, we thus amend the basic relationship so as to allow for a distance measure, as in the gravity literature. Our benchmark model specification thus becomes

$$\Delta c_{i,j}^t = \beta_0 + \beta_1 \Delta e_{i}^t + \beta_2 (\Delta e_{i}^t \ast \text{border}_{i,j}) + \beta_3 \Delta e_{i}^t \ast \ln d_{i,j} + v_{i,j}^t,$$

(2.3)

\[^{11}\text{Note that this equation is not literally a test of (2.2), since that specification implies that the relationship holds exactly. But extending the model to incomplete markets (as shown in Section 3 below) implies that the relationship between consumption and the real exchange rate holds only in conditional expectation. This can be used as more direct motivation for the regression specification here.}\]
where $\ln \tilde{d}_{i,j}$ is the normalized log distance between any two locations $i$ and $j$, defined as $\ln \tilde{d}_{i,j} = \ln d_{i,j} - \ln \overline{d}_{i,j}$. Here $d_{i,j}$ is the distance between locations $i$ and $j$, which we proxy using the distance between the capital cities of various jurisdictions; while $\ln \overline{d}_{i,j}$ is the average log distance between all cross-border pairs. This normalization implies that $\ln \tilde{d}_{i,j}$ is equal to zero at $\ln d_{i,j} = \ln \overline{d}_{i,j}$, and simplifies interpretation of the $\beta_2$ coefficient, which now expresses the average effect of the border for the consumption-real exchange rate relationship between any two locations that are $\ln \overline{d}_{i,j}$ kilometers away. The interaction term between the real exchange rate and distance allows the relationship between $\Delta c_{i}^{j,i}$ and $\Delta c_{i}^{j,i}$ to change monotonically with the distance. In all regressions standard errors are clustered at the country-pair level.

### Table 1: Estimates of Border Effect

<table>
<thead>
<tr>
<th></th>
<th>Pooled</th>
<th>Fixed effects</th>
<th>Pooled</th>
<th>Fixed effects</th>
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<tr>
<td></td>
<td>(i)</td>
<td>(ii)</td>
<td>(iii)</td>
<td>(iv)</td>
</tr>
<tr>
<td>$\Delta e_{i}^{j,i}$</td>
<td>0.349***</td>
<td>0.373***</td>
<td>0.275***</td>
<td>0.298***</td>
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<tr>
<td></td>
<td>(0.085)</td>
<td>(0.064)</td>
<td>(0.062)</td>
<td>(0.046)</td>
</tr>
<tr>
<td>$\Delta e_{i}^{j,i}$*$\text{border}_{i,j}$</td>
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<td>-0.385***</td>
<td>-0.291***</td>
<td>-0.305***</td>
</tr>
<tr>
<td></td>
<td>(0.081)</td>
<td>(0.063)</td>
<td>(0.060)</td>
<td>(0.049)</td>
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<tr>
<td>$\Delta e_{i}^{j,i}$*$\ln \tilde{d}_{i,j}$</td>
<td>0.021</td>
<td>0.026</td>
<td>0.019</td>
<td>0.022</td>
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<td></td>
<td>(0.013)</td>
<td>(0.020)</td>
<td>(0.013)</td>
<td>(0.019)</td>
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<tr>
<td>$\Delta y_{i}^{j,i}$</td>
<td>0.169***</td>
<td></td>
<td>0.161***</td>
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<tr>
<td></td>
<td>(0.024)</td>
<td></td>
<td>(0.027)</td>
<td></td>
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<tr>
<td>$\Delta y_{i}^{j,i}$*$\text{border}_{i,j}$</td>
<td>-0.036</td>
<td></td>
<td>-0.050</td>
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</tr>
<tr>
<td></td>
<td>(0.036)</td>
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<td>(0.038)</td>
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<tr>
<td>$\Delta y_{i}^{j,i}$*$\ln \tilde{d}_{i,j}$</td>
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<td></td>
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<td></td>
<td>(0.028)</td>
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<td>(0.031)</td>
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<tr>
<td>$\Delta c_{i}^{j,i}$ + $\Delta c_{i}^{j,i}$*$\text{border}_{i,j}$</td>
<td>-0.022</td>
<td>-0.012</td>
<td>-0.016</td>
<td>-0.008</td>
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<tr>
<td></td>
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<td>(0.016)</td>
<td>(0.013)</td>
<td>(0.017)</td>
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Notes: The dependent variable is relative consumption growth between locations $i$ and $j$, $\Delta c_{i}^{j,i}$. The estimated specification in columns (i)-(ii) is equation (2.3), while in columns (iii)-(iv) it is equation (2.4). Standard errors are clustered at the country-pair level and are reported in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively.

Our findings from the pooled OLS and fixed effects estimation of equation (2.3) are presented in columns (i) and (ii) of Table 1. The results in column (i) indicate that the conditional correlation between the growth rates of real exchange rate and relative consumption within our sample of coun-

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12 We measure distance in kilometers.

13 In the fixed effects regression the fixed effects capture the time-invariant, bilateral pair specific jurisdiction effects.
tries is positive and significant, equal to 0.35 on average. Hence, within countries, the risk-sharing relationship between relative consumption and real exchange rates (which are cross jurisdictional differences in rates of CPI inflation) holds quite strongly – regions with higher relative consumption growth exhibit depreciating relative cross-jurisdictional real exchange rates.

But when we interact real exchange rate changes with the border, the estimated border coefficient is negative. It is also large and economically significant. In fact, due to this effect, the consumption-real exchange rate correlation across countries turns negative, equal to \(-0.022\) on average, although this correlation is not significantly different from zero. Taking equation \((2.2)\) as our basic theory of risk-sharing, these results suggest that relative prices facilitate risk-sharing within countries, but impede risk-sharing across countries. The estimates in column (ii) obtained from the fixed effects regression confirm this finding.\(^{14}\)

Many studies of risk-sharing have relaxed the assumption of complete access to capital markets and posited the alternative specification in which at least a fraction of consumers do not make consumption plans based solely on intertemporal optimization, but may alternatively follow rules of thumb, or in fact have no ability to borrow and lend at all.\(^{15}\) To allow for this, we extend our framework to encompass limited capital market participation. Say that a fraction of households are hand-to-mouth consumers – that is they are restricted to consume only their current income. The testable implication of such a modified model is that relative consumption growth of these hand-to-mouth consumers living in any two locations follows their relative income growth. Let \(\Delta y_{i,j}^t = \Delta \ln Y_{i,t} - \Delta \ln Y_{j,t}\) denote the relative income growth between locations \(i\) and \(j\) at time \(t\). Then the relationship in equation \((2.3)\) must be modified to account for the limited participation

\(^{14}\)Note that the coefficient on the interactive term between RER and log distance indicates that the consumption-RER correlation increases slightly for the jurisdictions that are located further away from each other.

as follows:

$$
\Delta c_i^{j,t} = \beta_0 + \beta_1 \Delta c_i^{j,t} + \beta_2 (\Delta c_i^{j,t} \times \text{border}_{i,j}) + \beta_3 \Delta c_i^{j,t} \ln d_{i,j} \\
+ \beta_4 \Delta y_i^{j,t} + \beta_5 (\Delta y_i^{j,t} \times \text{border}_{i,j}) + \beta_6 \Delta y_i^{j,t} \ln d_{i,j} + v_i^{j,t} 
$$  
(2.4)

This specification allows for the border to affect the consumption-income relationship, and also for the relationship between $\Delta c_i^{j,t}$ and $\Delta y_i^{j,t}$ to change monotonically with distance. The results from this estimation are presented in columns (iii) and (iv) of Table 1. We find that there is significant positive association between $\Delta c_i^{j,t}$ and $\Delta y_i^{j,t}$ in both pooled and fixed effects specifications. Allowing for limited participation also affects the within-country correlation between $\Delta c_i^{j,t}$ and $\Delta e_i^{j,t}$ as it declines to about 0.28. At the same time, the border effect in the consumption-real exchange rate relationship remains negative and significant. As before, with the addition of the border coefficient, the relationship between $\Delta c_i^{j,t}$ and $\Delta e_i^{j,t}$ becomes slightly negative. Overall, our estimated border effect in the consumption-real exchange rate risk-sharing remains robust to possibility of limited participation in domestic and international capital markets.

### 2.3 What drives the border effect?

An obvious reason for why consumption-real exchange rate correlation drops once cross-border pairs are considered is that within-country real exchange rates compare only relative inflation differentials between regions, while across-country real exchange rates involve comparisons across currencies and so involve nominal exchange rate changes. For this purpose we decompose the real exchange rate into its components as follows. Recall the definition of the real exchange rate:

$$
RER_t^{j,i} = P_{j,t}S_t^{i,j} / P_{i,t}.
$$

Taking logs and first-differencing we get

$$
\Delta \ln RER_t^{j,i} = \Delta \ln (P_{j,t} / P_{i,t}) + \Delta \ln S_t^{i,j}, 
$$  
(2.5)
where the first term on the right-hand-side captures movements in the real exchange rate due to local rates of goods price inflation, while the second term is due to the movements in the nominal exchange rate. Using this decomposition, we amend the specification in equations (2.3) and (2.4) to include the growth rate in nominal exchange rates. The results from fixed effects regressions are presented in columns (i) and (ii) of Table 2.

Table 2: Estimates of the Border Effect: RER Decomposition

<table>
<thead>
<tr>
<th></th>
<th>(i)</th>
<th>(ii)</th>
<th>(iii)</th>
<th>(iv)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta c_{t}^{i,j}$</td>
<td>0.384***</td>
<td>0.308***</td>
<td>0.387***</td>
<td>0.311***</td>
</tr>
<tr>
<td></td>
<td>(0.060)</td>
<td>(0.043)</td>
<td>(0.059)</td>
<td>(0.042)</td>
</tr>
<tr>
<td>$\Delta c_{t}^{i,j} \times \text{border}_{i,j}$</td>
<td>-0.127</td>
<td>-0.061</td>
<td>-0.137</td>
<td>-0.068</td>
</tr>
<tr>
<td></td>
<td>(0.122)</td>
<td>(0.082)</td>
<td>(0.123)</td>
<td>(0.083)</td>
</tr>
<tr>
<td>$\Delta c_{t}^{i,j} \times \ln \delta_{i,j}$</td>
<td>0.038*</td>
<td>0.034</td>
<td>0.042**</td>
<td>0.037*</td>
</tr>
<tr>
<td></td>
<td>(0.020)</td>
<td>(0.021)</td>
<td>(0.020)</td>
<td>(0.020)</td>
</tr>
<tr>
<td>$\Delta g_{t}^{i,j}$</td>
<td>0.159**</td>
<td>0.160***</td>
<td>0.159**</td>
<td>0.160***</td>
</tr>
<tr>
<td></td>
<td>(0.026)</td>
<td>(0.026)</td>
<td>(0.026)</td>
<td>(0.026)</td>
</tr>
<tr>
<td>$\Delta y_{t}^{i,j} \times \text{border}_{i,j}$</td>
<td>-0.049</td>
<td>-0.052</td>
<td>-0.052</td>
<td>-0.052</td>
</tr>
<tr>
<td></td>
<td>(0.038)</td>
<td>(0.040)</td>
<td>(0.040)</td>
<td>(0.040)</td>
</tr>
<tr>
<td>$\Delta y_{t}^{i,j} \times \ln \delta_{i,j}$</td>
<td>0.037</td>
<td>0.038</td>
<td>0.038</td>
<td>0.038</td>
</tr>
<tr>
<td></td>
<td>(0.029)</td>
<td>(0.030)</td>
<td>(0.030)</td>
<td>(0.030)</td>
</tr>
<tr>
<td>$\Delta \ln S_{t}^{i,j}$</td>
<td>-0.293**</td>
<td>-0.278***</td>
<td>-0.290**</td>
<td>-0.276***</td>
</tr>
<tr>
<td></td>
<td>(0.110)</td>
<td>(0.076)</td>
<td>(0.111)</td>
<td>(0.077)</td>
</tr>
<tr>
<td>$\Delta \Delta e_{t}^{i,j} \times \text{border}_{i,j}^{SPA-GER}$</td>
<td>0.481***</td>
<td>0.375***</td>
<td>0.481***</td>
<td>0.375***</td>
</tr>
<tr>
<td></td>
<td>(0.043)</td>
<td>(0.048)</td>
<td>(0.043)</td>
<td>(0.048)</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>(i)</th>
<th>(ii)</th>
<th>(iii)</th>
<th>(iv)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta c_{t}^{i,j} + \Delta e_{t}^{i,j} \times \text{border}_{i,j}$</td>
<td>0.257**</td>
<td>0.248***</td>
<td>0.250**</td>
<td>0.242***</td>
</tr>
<tr>
<td></td>
<td>(0.106)</td>
<td>(0.072)</td>
<td>(0.107)</td>
<td>(0.073)</td>
</tr>
<tr>
<td>$\Delta c_{t}^{i,j} + \Delta e_{t}^{i,j} \times \text{border}<em>{i,j} + \Delta \ln S</em>{t}^{i,j}$</td>
<td>-0.030*</td>
<td>-0.031</td>
<td>-0.039**</td>
<td>-0.033*</td>
</tr>
<tr>
<td></td>
<td>(0.019)</td>
<td>(0.020)</td>
<td>(0.018)</td>
<td>(0.019)</td>
</tr>
<tr>
<td>$\Delta c_{t}^{i,j} + \Delta e_{t}^{i,j} \times \text{border}<em>{i,j} + \Delta \ln S</em>{t}^{i,j} + \Delta c_{t}^{i,j} \times \text{border}_{i,j}^{SPA-GER}$</td>
<td>0.441***</td>
<td>0.342***</td>
<td>0.441***</td>
<td>0.342***</td>
</tr>
<tr>
<td></td>
<td>(0.041)</td>
<td>(0.052)</td>
<td>(0.041)</td>
<td>(0.052)</td>
</tr>
</tbody>
</table>

Notes: The dependent variable is relative consumption growth between locations $i$ and $j$, $c_{t}^{i,j}$. The estimated specification in columns (i) and (iii) is equation (2.3), while in column (ii) and (iv) is equation (2.4). All are modified to include the growth rate of the nominal exchange rate. All regressions are estimated with bilateral jurisdiction pair-specific fixed effects. Standard errors are clustered at the country-pair level and are reported in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively.

Columns (i) provides estimates of specification (2.3), while column (ii) allows for market segmentation and thus summarizes the estimates of specification (2.4), both amended to include the nominal exchange rate growth rate between jurisdictions $i$ and $j$. The key result that stands out from Table 2 is the coefficient on the nominal exchange rate growth is negative and significant. Furthermore, the negative coefficient on the border effect is substantially reduced when we control for nominal exchange rate movements. This suggests that a sizeable component of the border effect
from Table I can be accounted for by nominal exchange rate movements. Thus, without at this stage suggesting causation, the finding indicates that country pairs with higher nominal exchange rate volatility will deviate more from the benchmark model of full risk sharing.

Our results do not imply that the entire border effect is accounted for by the nominal exchange rate. The negative effect of the border is reduced when the nominal exchange rate is added to the regression specification, but it is not eliminated. There are many possible reasons for why the border effect may remain important. Risk sharing within countries may be greater than that across countries due to deeper financial integration; common language, culture and institutions; lower formal and informal trade barriers; greater labor and capital mobility; presence of fiscal insurance mechanisms; greater synchronization of shock processes within countries; differences in demand elasticities across countries (pricing-to-market); etc. We investigate the quantitative role played by some of these factors below, using the model. Our results here, however, point to the nominal exchange rate as an important new driver of the border effect in risk-sharing.

2.4 Disentangling the border from the nominal exchange rate

The results show that the consumption-real exchange rate correlation drops significantly when comparing across borders, and that the nominal exchange rate accounts for a significant part of this decrease. This does not, however, mean that eliminating nominal exchange rate flexibility would lead to an increase in the consumption-real exchange rate correlation. The reason is that all our cross-border comparisons also involve nominal exchange rates. If there are attributes of cross-border relationships that are correlated with the nominal exchange rate, we may mistakenly ascribe the significant fall in the consumption-real exchange rate correlation to the nominal exchange rate, while it is in fact independent of exchange rate policy.
2.4.1 Spain-Germany case

One way to disentangle the effects of the nominal exchange rate from the border is by focusing on cross-border pairs within the sample that exhibit little nominal exchange rate volatility. Our sample includes Germany and Spain. These two countries have had either very low or zero bilateral nominal exchange rate variability during our sample period. To further isolate the role of the nominal exchange rate, we can therefore compare the properties of all regional pairs across the Spain-Germany border with the cross-border pairs whose exchange rate is freely floating.

Table 3 shows the mean standard deviation of monthly exchange rate changes between all cross-border pairs of countries for our sample.

<table>
<thead>
<tr>
<th></th>
<th>US</th>
<th>Japan</th>
<th>Germany</th>
<th>Spain</th>
</tr>
</thead>
<tbody>
<tr>
<td>Canada</td>
<td>1.3</td>
<td>2.9</td>
<td>2.6</td>
<td>2.6</td>
</tr>
<tr>
<td>US</td>
<td>-</td>
<td>2.7</td>
<td>2.3</td>
<td>2.3</td>
</tr>
<tr>
<td>Japan</td>
<td>-</td>
<td>-</td>
<td>2.9</td>
<td>3.1</td>
</tr>
<tr>
<td>Germany</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td>0.6</td>
</tr>
</tbody>
</table>

Notes: Standard deviations of changes in monthly bilateral nominal exchange rates over samples matching those in our bilateral consumption and real exchange rate dataset.

It is apparent that the outlier is Germany and Spain. The cross-border sample for Germany and Spain covers the period from 1995 to 2004. For the first four years, Germany and Spain had separate currencies, but very low bilateral exchange rate movements. From 1999 onward, they had zero bilateral nominal exchange rate movements. Over the whole sample, the mean monthly standard deviation of nominal exchange rate changes is only 0.6%. This contrasts with much higher nominal exchange rate volatility for all other cross-border pairs. If it is the nominal exchange rate itself rather than some aspect of the border (which may be correlated with nominal exchange rates) that is causing the drop in the consumption-real exchange rate correlation for across-country pairs, we should expect that including a separate control for Germany-Spain comparisons in the
basic regressions would reverse some of the negative border effect. With this in mind, we extend
the specifications in columns (i) and (ii) of Table 2 to include an interactive term between $\Delta e_{ij}^{t}$
and a Germany-Spain border dummy that takes a value of one for all jurisdiction pairs located
across the border between Germany and Spain, and zero otherwise. The results are presented
in column (iii) of Table 2 for the benchmark specification, and in column (iv) for the specification
with hand-to-mouth consumers. We find that the coefficient on $\Delta e_{ij}^{t} \cdot \text{border}_{i,j}^{G\text{ER}-S\text{PA}}$ is positive
and significant, and, importantly, it is large enough to turn the border effect between Germany
and Spain positive, equal to 0.44 in the benchmark specification and 0.34 in the specification with
hand-to-mouth consumers. This is in sharp contrast to the average border effect in the full sample
of countries equal to -0.039 in the benchmark specification and -0.033 in the specification with
hand-to-mouth consumers.

2.4.2 Cross-country evidence

To further assess the role of the exchange rate in consumption-real exchange rate correlations, we
turn to cross-country evidence. In particular, we study a cohort of the 12 European Union member
states during 1974-2007 period. To evaluate the role of the exchange rate, we split the sample
into pre-euro period (1974-1998) and euro period (1999-2007). For each subsample we compute
all bilateral correlations of relative consumption growth with the growth rate in the real exchange
rate for each country. The average correlations for each country and each sub-period are reported

---

16 As noted, since the nominal exchange rate between Germany and Spain is fixed from 2000 onwards we cannot
identify the contribution of the nominal exchange rate to the border effect between these two countries. Instead, we
estimate how the border effect between Germany and Spain differs from the average border effect estimated in the
full sample of countries.

17 As an additional check, we allowed for interaction term between the real exchange rate and the border for Spain
paired with all other countries (but Germany) and Germany paired with all other countries (but Spain) in our sample.
The estimates of the border effect for these pairs are negative and significant, similar in magnitude to the average
border effect. This suggests that our findings for the Germany-Spain pairs are not driven by factors specific to these
two countries, but rather by the attributes specific to Spain-Germany bilateral pairs. See Table S1 in the online
appendix for detailed results.

18 Data details and sources are provided in the Appendix A.1.

19 Greece adopted euro on January 1, 2001. We adjust our calculations accordingly.
in Table 4. A clear contrast arises between the two periods: In the pre-euro period, the relative consumption-real exchange rate correlation is negative, on average, across countries, equal to -0.19. The same correlation turns positive in the euro period and is equal to 0.05. The difference in correlations across the two sub-periods is also statistically significant at 1% level.

Table 4: International C-RER correlation: Fixed vs flexible regime

<table>
<thead>
<tr>
<th></th>
<th>Pre-euro period</th>
<th>Euro period</th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>-0.18</td>
<td>0.17</td>
</tr>
<tr>
<td>Belgium</td>
<td>-0.06</td>
<td>0.21</td>
</tr>
<tr>
<td>Finland</td>
<td>-0.49</td>
<td>0.48</td>
</tr>
<tr>
<td>France</td>
<td>-0.11</td>
<td>-0.05</td>
</tr>
<tr>
<td>Germany</td>
<td>-0.13</td>
<td>0.15</td>
</tr>
<tr>
<td>Greece</td>
<td>0.04</td>
<td>0.15</td>
</tr>
<tr>
<td>Ireland</td>
<td>-0.16</td>
<td>-0.24</td>
</tr>
<tr>
<td>Italy</td>
<td>-0.30</td>
<td>-0.13</td>
</tr>
<tr>
<td>Luxembourg</td>
<td>-0.07</td>
<td>0.00</td>
</tr>
<tr>
<td>Netherlands</td>
<td>-0.21</td>
<td>-0.24</td>
</tr>
<tr>
<td>Portugal</td>
<td>-0.45</td>
<td>0.14</td>
</tr>
<tr>
<td>Spain</td>
<td>-0.16</td>
<td>0.01</td>
</tr>
<tr>
<td>Average</td>
<td>-0.19</td>
<td>0.05</td>
</tr>
</tbody>
</table>

Overall, these results support our earlier findings on the role of exchange rate fluctuations for the finding of negative correlation between relative consumption growth and real exchange rate growth. Furthermore, our results in this subsection are in line with the findings in Hadzi-Vaskov (2008).

To summarize, we find that the relative price movements facilitate consumption risk-sharing within countries; while they obstruct consumption risk-sharing across country borders. Importantly, most of this border effect can be attributed to nominal exchange rate variability. In Table 2, the inclusion of the bilateral nominal exchange rate reduces the coefficient on the border by 67% and 80% in the case with and without hand-to-mouth consumers, respectively.

---

20 The test of significance is obtained using Fisher’s variance stabilizing transformation of correlation coefficients.
21 We recognize that besides the exchange rate regime, other factors could have also contributed to the increase in the consumption-RER correlation in the second sub-period. For instance, the degree of financial integration could have risen, or policy coordination could have increased in the latter period between the EU members. However, the evidence in this subsection, together with intra-national comparisons suggest to us that the exchange rate is an important part of the story.
22 Our analysis in this subsection differs from Hadzi-Vaskov (2008) in that we use annual data and an extra year of the data.
In the next sub-section we show the robustness of our findings using alternative regression specifications and different decomposition approaches.

2.5 Robustness

2.5.1 Country heterogeneity

In the regression specifications (2.3) and (2.4) the effect of the border on the consumption-real exchange rate correlation as captured by the interactive term on the real exchange rate with the border dummy, is measured relative to the average correlation in the intra-country pairs. Gorodnichenko and Tesar (2009) argue that in the presence of cross-country heterogeneity in the distribution of within-country price differentials, such an average benchmark is arbitrary and can lead to misleading results. In particular, in this case, the border coefficient may capture the joint effect of the border and country heterogeneity in the consumption-real exchange rate correlation.

We can address this issue by allowing the intra-national correlation between $\Delta c_{i,j}^t$ and $\Delta e_{i,j}^t$ to be country-specific. In particular, we amend the regression specification in equation (2.4) with the product terms of $\Delta e_{i,j}^t$ with a set of dummy variables, each identifying jurisdiction pairs within each country in our sample. In all cases we use within US states pairs as the benchmark. The coefficient on this interactive term will give us the change in the intra-national consumption-real exchange rate correlation when transitioning from US state pairs to other country pairs. The coefficient on the interactive term between $\Delta c_{i,j}^t$ and the border dummy, as before, will give us the effect of the border on the correlation, except now we can estimate the effect of the border crossing from the perspective of each individual country in our sample.

Table 5 reports our findings.\textsuperscript{23} For completeness, panel (a) summarizes our earlier results from the fixed effects regressions where we used the average correlation in the intra-country pairs as

\textsuperscript{23}Detailed estimation results are available from the authors upon request.
Table 5: Estimates of the Border Effect: Accounting for country heterogeneity

<table>
<thead>
<tr>
<th></th>
<th>no $\Delta y_{t}^{i,j}$</th>
<th>with $\Delta y_{t}^{i,j}$</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Panel (a). Homogeneous benchmark</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>Intra-national</strong></td>
<td>0.373***</td>
<td>0.298***</td>
</tr>
<tr>
<td></td>
<td>(0.064)</td>
<td>(0.046)</td>
</tr>
<tr>
<td><strong>International</strong></td>
<td>-0.012</td>
<td>-0.008</td>
</tr>
<tr>
<td></td>
<td>(0.016)</td>
<td>(0.017)</td>
</tr>
<tr>
<td><strong>Panel (b). Heterogeneous benchmarks</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>Intra-national correlations</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>US</td>
<td>0.433***</td>
<td>0.329***</td>
</tr>
<tr>
<td></td>
<td>(0.011)</td>
<td>(0.007)</td>
</tr>
<tr>
<td>Canada</td>
<td>0.214***</td>
<td>0.159***</td>
</tr>
<tr>
<td></td>
<td>(0.009)</td>
<td>(0.006)</td>
</tr>
<tr>
<td>Germany</td>
<td>-0.201***</td>
<td>0.135**</td>
</tr>
<tr>
<td></td>
<td>(0.048)</td>
<td>(0.050)</td>
</tr>
<tr>
<td>Japan</td>
<td>0.360***</td>
<td>0.227***</td>
</tr>
<tr>
<td></td>
<td>(0.043)</td>
<td>(0.042)</td>
</tr>
<tr>
<td>Spain</td>
<td>0.925***</td>
<td>1.151***</td>
</tr>
<tr>
<td></td>
<td>(0.032)</td>
<td>(0.053)</td>
</tr>
<tr>
<td><strong>International correlations</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>US</td>
<td>-0.007</td>
<td>-0.005</td>
</tr>
<tr>
<td></td>
<td>(0.017)</td>
<td>(0.019)</td>
</tr>
<tr>
<td>Canada</td>
<td>-0.226***</td>
<td>-0.175***</td>
</tr>
<tr>
<td></td>
<td>(0.019)</td>
<td>(0.022)</td>
</tr>
<tr>
<td>Germany</td>
<td>-0.641***</td>
<td>-0.199***</td>
</tr>
<tr>
<td></td>
<td>(0.029)</td>
<td>(0.033)</td>
</tr>
<tr>
<td>Japan</td>
<td>-0.080***</td>
<td>-0.107***</td>
</tr>
<tr>
<td></td>
<td>(0.025)</td>
<td>(0.025)</td>
</tr>
<tr>
<td>Spain</td>
<td>0.485***</td>
<td>0.817***</td>
</tr>
<tr>
<td></td>
<td>(0.017)</td>
<td>(0.041)</td>
</tr>
</tbody>
</table>

Notes: The dependent variable is relative consumption growth between locations $i$ and $j$, $\Delta c_{t}^{i,j}$. Panel (a) summarizes the results from columns (ii) and (iv) in Table 1. Panel (b) reports the results from a regression specification in equation (2.3) – column labelled "no $\Delta y_{t}^{i,j}$", and specification in equation (2.4) – column labelled "with $\Delta y_{t}^{i,j}$", both modified to include product terms between $\Delta c_{t}^{i,j}$ and dummies that identify all within country pairs for Canada, Germany, Japan and Spain. Within US state pairs are used as the benchmark. All regressions include jurisdiction pairs fixed effects. Standard errors are clustered at the country-pair level and are reported in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively.

Our results reveal some amount of heterogeneity across our sample of countries. However, some common patterns emerge. First, the intra-national consumption-real exchange rate correlations are consistently positive, with the US and Spain showing the highest numbers in our sample, while...
Germany is characterized by the lowest correlation. Second, international correlations are negative for all countries, except Spain. Note however that even for Spain, international correlations are still considerably lower than national correlations. The correlations are the lowest (most negative) in Germany.

2.5.2 Other controls

One potential criticism of our benchmark specification derived in equation (2.2) is the fact that it assumes a separable utility function. A number of variables can potentially affect the marginal utility of consumption, such as leisure, government consumption, real money balances, etc. Ravn (2001) investigates the importance of these variables for the consumption-real exchange rate correlation in cross-country data. He finds that the negative correlation between relative consumption and the real exchange rate remains robust to relaxing the separability assumption. Unfortunately, a similar exercise can not be conducted with our data because intra-jurisdiction information for these variables is not available.

One way to account for the omitted variables for which data or proxies are not available is to use a lagged dependent variable (i.e. relative consumption growth) in specifications (2.3)-(2.4). Such an extension would also arise from a model with exogenous habit persistence, as in Abel (1990). In dynamic panels, however, the unobserved panel-level effects are correlated with the lagged dependent variables, making standard estimators inconsistent. To obtain consistent estimates for model parameters we use Arellano and Bond (1991) generalized method of moments (GMM) approach. The method uses the first difference of all the exogenous variables and second and higher lags of the dependent variable and all other variables as instruments. Standard errors are clustered at the country-pair level. The results from the estimation are presented in Table 6.

25The only exception is Germany in the specification with no hand-to-mouth consumers, who exhibits negative intra-national consumption-RER correlation.
Table 6: Estimates of Border Effect: Lagged consumption growth

<table>
<thead>
<tr>
<th></th>
<th>(i)</th>
<th>(ii)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta e_{i}^{ij}$</td>
<td>0.161***</td>
<td>0.062**</td>
</tr>
<tr>
<td></td>
<td>(0.043)</td>
<td>(0.032)</td>
</tr>
<tr>
<td>$\Delta e_{i}^{ij} \times \text{border}_{i,j}$</td>
<td>-0.207***</td>
<td>-0.094***</td>
</tr>
<tr>
<td></td>
<td>(0.060)</td>
<td>(0.038)</td>
</tr>
<tr>
<td>$\Delta e_{i}^{ij} \times \ln \tilde{d}_{i,j}$</td>
<td>0.063***</td>
<td>0.053***</td>
</tr>
<tr>
<td></td>
<td>(0.018)</td>
<td>(0.017)</td>
</tr>
<tr>
<td>$\Delta y_{i}^{ij}$</td>
<td></td>
<td>0.184***</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.028)</td>
</tr>
<tr>
<td>$\Delta y_{i}^{ij} \times \text{border}_{i,j}$</td>
<td>0.120</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.091)</td>
</tr>
<tr>
<td>$\Delta y_{i}^{ij} \times \ln \tilde{d}_{i,j}$</td>
<td>0.081</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.091)</td>
</tr>
<tr>
<td>$\Delta \tilde{d}_{i-1}^{ij}$</td>
<td>0.013</td>
<td>-0.007</td>
</tr>
<tr>
<td></td>
<td>(0.010)</td>
<td>(0.015)</td>
</tr>
<tr>
<td>$\Delta e_{i}^{ij} + \Delta e_{i}^{ij} \times \text{border}_{i,j}$</td>
<td>-0.046**</td>
<td>-0.032</td>
</tr>
<tr>
<td></td>
<td>(0.023)</td>
<td>(0.021)</td>
</tr>
</tbody>
</table>

N: 124588 124588

Notes: The dependent variable is relative consumption growth between locations $i$ and $j$, $\Delta e_{i}^{ij}$. The estimated specification in columns (i) is equation (2.3), while in columns (ii) it is equation (2.4). Each regression also includes one lag of the dependent variable. All equations are estimated with Arellano and Bond (1991) GMM method. Standard errors are clustered at the country-pair level and are reported in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively.

The lagged consumption variable is negative and generally statistically significant. At the same time, our results for the consumption-real exchange rate correlation remain largely robust. If anything, the introduction of a lagged relative consumption growth into our estimation, worsens the international evidence on consumption-real exchange rate risk-sharing by producing a more negative correlation between the two variables. The intra-national correlation remains positive and significant. Importantly, our results also remain robust when we include the nominal exchange rate explicitly in the regressions. The nominal exchange rate still accounts for a large share of the border effect in risk-sharing.26

2.5.3 An alternative decomposition and the role of the nominal exchange rate

Up to now, we have assumed that the benchmark for efficient risk-sharing in the data is that relative consumption growth rates across any two locations should be positively associated with the real exchange rate. However, in this section, we explore an alternative decomposition of the border effect and the role of the nominal exchange rate. These results can be found in the online appendix to the paper.
exchange rate, independent of other variables. But these tests do not give a clear metric for the degree to which risk-sharing fails in the data, nor the extent to which the border and the nominal exchange rate contribute to this failure. Here we follow in the spirit of Engel and Rogers (1996) in constructing a welfare-relevant measure of the failure of risk-sharing. We then ask how much the border and the nominal exchange rate can account for this measure.\footnote{Engel and Rogers (1996) measure the extent of the failure of the law of one price using the standard deviation of the price differentials of similar goods across cities in the US and Canada.}

We return to equation (2.2) and for each jurisdiction pair in our sample, for a choice of $\sigma$, we compute $\Delta c_{i,j}^i - (1/\sigma)\Delta e_{i,j}^j$ and obtain its standard deviation in the time-series.\footnote{We also conduct the analysis of this section using the mean of the squared difference, $\text{mean} \left[ \Delta c_{i,j}^i - (1/\sigma)\Delta e_{i,j}^j \right]^2$, and find that the results are robust to this alternative measure of risk-sharing.} This gives us a cross-section of such standard deviations. If our choice of $\sigma$ is correct, then with full risk-sharing this statistic would be zero for each bilateral pair. We set $\sigma = 2$, so that the elasticity of intertemporal substitution is equal to 0.5 — a standard value in the literature. The summary statistics of this measure are reported in row (i) of Table 7.

<table>
<thead>
<tr>
<th>Variable</th>
<th>Obs</th>
<th>Mean</th>
<th>Std. Dev.</th>
<th>Min</th>
<th>Max</th>
</tr>
</thead>
<tbody>
<tr>
<td>(i) std.dev. $\left[ \Delta c_{i,j}^i - (1/\sigma)\Delta e_{i,j}^j \right]$</td>
<td>10146</td>
<td>0.0514</td>
<td>0.0183</td>
<td>0.0039</td>
<td>0.1391</td>
</tr>
<tr>
<td>(ii) std.dev. $\left[ \Delta c_{i,j}^i - (1/\sigma)\Delta e_{i,j}^j \right]$</td>
<td>10146</td>
<td>0.0332</td>
<td>0.0158</td>
<td>0.0039</td>
<td>0.1095</td>
</tr>
</tbody>
</table>

Notes: The table reports summary statistics of the presented variables. Obs. refer to the number of observations in each sample; Mean - sample average; Std. Dev. - sample standard deviation; Min-sample minimum; Max-sample maximum.

Using this measure of the departure from full risk-sharing, we estimate the following regression:

$$\text{std.dev.} \left[ \Delta c_{i,j}^i - (1/\sigma)\Delta e_{i,j}^j \right] = \beta_1 \text{border}_{i,j} + \sum_{m=1}^{5} \alpha_m D_{mm} + \varepsilon_{i,j}, \quad (2.6)$$

where $\text{border}_{i,j}$, is as before a dummy variable that takes on a value of one if locations $i$ and $j$ are in separate countries, and zero otherwise, while $\varepsilon_{i,j}$ is the regression error term. We also consider a specification in which we include the log of the distance between locations $i$ and $j$. To
account for potential heterogeneity across countries in the average volatility of deviations from perfect risk-sharing, we also include a set of dummies, identifying within-country pairs, \( D_{mm} \).\(^{29}\)

Table 8 reports the results. Panel (i) presents the results without the log distance variable in the specification, while panel (ii) - includes the log distance variable. The effect of the border is positive and significant. It means that controlling for country-specific volatility in the risk-sharing measure, there is a greater deviation from efficient risk-sharing when comparing across countries relative to comparing within countries. This result holds whether we control for distance in our estimation or not. The average volatility of our risk-sharing measure is 5.14%, while the border coefficient is 2.74%, so the border accounts for 53% of the total, controlling for distance. Without the distance control, the contribution of the border is even larger at 77%.

Table 8: Alternative measure of risk-sharing: The role of the border

<table>
<thead>
<tr>
<th>Variable</th>
<th>border ( \Delta \sigma_{i;j} - (1/\sigma)\Delta \sigma_{i}^2 )</th>
<th>( \ln(\text{distance}) )</th>
<th>( N )</th>
<th>explained by border</th>
</tr>
</thead>
<tbody>
<tr>
<td>(i) std.dev. ( \Delta \sigma_{i;j} - (1/\sigma)\Delta \sigma_{i}^2 )</td>
<td>0.0395***</td>
<td>10146</td>
<td>0.77</td>
<td></td>
</tr>
<tr>
<td>(ii) std.dev. ( \Delta \sigma_{i;j} - (1/\sigma)\Delta \sigma_{i}^2 )</td>
<td>0.0274***</td>
<td>0.0090**</td>
<td>10146</td>
<td>0.53</td>
</tr>
<tr>
<td>(iii) std.dev. ( \Delta \sigma_{i;j} - (1/\sigma)\Delta \sigma_{i}^2 )</td>
<td>0.0149***</td>
<td>(0.0064)</td>
<td>(0.0038)</td>
<td>0.45</td>
</tr>
<tr>
<td>(iv) std.dev. ( \Delta \sigma_{i;j} - (1/\sigma)\Delta \sigma_{i}^2 )</td>
<td>0.0113**</td>
<td>0.0027</td>
<td>10146</td>
<td>0.34</td>
</tr>
</tbody>
</table>

Notes: This table presents results from estimating regression specification (2.6). The dependent variable is noted in the first column of the table. Standard errors are clustered at the country-pair level and are reported in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively.

To assess the extent to which this border effect be attributed to movements in the nominal exchange rate we decompose the real exchange rate into that component driven by relative price levels and that by the nominal exchange rate.\(^{30}\)

\[
\Delta \ln RER_{t}^{i;j} = \Delta \ln(P_{j,t}S_{i,j}/P_{i,t}) + \Delta \ln(S_{t}^{i;j}/S_{i;j}), \quad (2.7)
\]

\(^{29}\)Note that this is a cross-sectional regression.

\(^{30}\)In this example, the test is purely a cross-section one, so we cannot include nominal exchange rate variability as a right-hand side variable.
where $\bar{S}_{i,j}$ represents the sample average nominal exchange rate for bilateral pair $i,j$. To get a measure of the importance of nominal exchange rate volatility in this measure of risk-sharing, we construct a proxy real exchange rate $\widetilde{RER}_{j,i}$, where

$$\widetilde{RER}_{j,i} \equiv \ln(P_{j,t} \bar{S}_{i,j}/P_{i,t})$$

This gives a measure of the real exchange rate when there is no nominal exchange rate variability so that the real exchange rate reflects only movements in relative prices, as in the case of within-country variability. This decomposition parallels the decomposition between relative prices and nominal exchange rate in our panel regressions. After computing $\widetilde{RER}_{j,i}$, we repeat the steps above to construct the volatility of our risk-sharing measure for every jurisdiction pair as $\text{std.dev.} \left[ \Delta c_{t}^{j,i} - (1/\sigma)\Delta \tilde{c}_{t}^{j,i} \right]$, where $\Delta \tilde{c}_{t}^{j,i} = \ln(\widetilde{RER}_{t}^{j,i}) - \ln(\widetilde{RER}_{t-1}^{j,i})$. Row (ii) of Table 7 presents the summary statistics on this adjusted measure. Clearly, this measure is less volatile than the original measure based on unadjusted real exchange rate. We then estimate the regression specification in equation (2.6) using this adjusted measure of deviations from the perfect risk-sharing. The results are reported in panels (iii) and (iv) of Table 8. We find that the coefficient on the border dummy is still positive and significant. But it is substantially smaller than before. This result is consistent with our findings from the decomposition approach in Section 2.3.

Specifically, from panel (iv), the coefficient on the border dummy is 1.1%, while the average volatility of the adjusted risk-sharing measure as reported in Table 7 is 3.3%. So the border accounts

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31 Engel and Rogers (1996) conduct a similar decomposition to isolate the effects of the nominal exchange rate on deviations from the law of one price across countries.

32 We note that interpreting the risk-sharing measure using $\widetilde{RER}_{t}$ rather than $RER_{t}$ implicitly assumes that relative nominal prices when compared across countries would display the same characteristics with a fixed exchange rate as they have done historically under a floating exchange rate. While it is well known that relative price levels are an order of magnitude less volatile than nominal exchange rates in almost all low-inflation countries (e.g. Engel (1999)), we do not have any way to directly test this assumption. Nevertheless, the approach here is helpful in the sense that it offers a direct estimate of the importance of nominal exchange rate variation in deviations from risk-sharing that can be compared to the estimates implied by the panel regressions.
for 34% of the standard deviation of this risk-sharing measure. Thus, when the nominal exchange rate is dropped from the calculation of the relative prices, the contribution of the border to the average volatility of the risk-sharing measure declines from 53% to 34%, controlling for distance. Thus, based on this measure, more than one-third of the border effect can be attributed to nominal exchange rate variation. The effect of the nominal exchange rate variation on the border is even larger without the distance control, equal to 42%. These results suggest a substantial role for the nominal exchange rate in accounting for deviations from full cross country risk-sharing.

3 Consumption Risk-Sharing with Sticky Prices

We have shown that the exchange rate plays a key role in empirical tests of the relationship between bilateral consumption differences and real exchange rate changes. Within countries, relative consumption and real exchange rates are generally positively correlated. Across countries, the same results seems to apply, in the absence of nominal exchange rate adjustment. But across countries with nominal exchange rate volatility, the correlation is much lower, and in fact is negative for the full sample. It seems then that the nominal exchange rate regime, or the degree of exchange rate flexibility, has consequences for cross-country risk-sharing.

This result cannot be easily rationalized by the existing structural models of risk-sharing. In principle, if goods prices are sticky, the behavior of the real exchange rate and the correlation of relative consumption and the real exchange rate may differ across exchange rate regimes. But this is not an automatic implication - many sticky price models (e.g. Clarida, Gali, and Gertler 2002, Devereux and Engel 2003) display volatile real and nominal exchange rates, but still have the property that the cross country risk sharing condition between consumption and real exchange rates holds exactly. Conversely, many potential resolutions to the Backus-Smith puzzle have no role at all for the nominal exchange rate, and so cannot offer a robust explanation of the findings.
of the previous section (since fixing the exchange in these models would have no implications for the empirical tests of risk-sharing).

In this section, we offer an accounting of the consumption-real exchange rate correlation that is tied directly to the nominal exchange rate. We combine features of the previous literature on the Backus-Smith puzzle with a two country, two region ‘New Keynesian’ model with gradual price adjustment. We then ask whether this model exhibits the property that the within-country consumption real exchange rate correlation is positive, and that the sign of the cross-country consumption-real exchange rate correlation depends upon the exchange rate regime.

The model is essentially an extension of the literature on international real business cycles, in particular the model of Stockman and Tesar (1995), to an environment with multiple regions within countries, and with sticky prices and alternative exchange rate regimes. Within countries, there is no exchange rate adjustment and our calibrated model predicts a positive correlation. Across countries, productivity shocks can result in the nominal and real exchange rate and relative consumption moving in different directions. By contrast, under a fixed exchange rate, the effect of the same shocks on the cross-country real exchange rate is limited. In our baseline model, we show that this difference can be enough to reverse the sign of the international consumption-real exchange rate correlation between exchange rate regimes. This result is especially highlighted when price adjustment is more backward looking.

3.1 A Model of the Consumption-Real Exchange Rate Correlation

The structure of the model is similar to that of Stockman and Tesar (1995) and Benigno and Thoenissen (2008). There are two countries, home and foreign. We also add a regional dimension within each country, so in both the home and foreign countries there are two regions, east and west. We assume that within each region there is a non-traded good, and a traded good produced and
consumed domestically, and consumed by other regions in the domestic country and the foreign
country. Hence, traded goods aggregate consist of goods from both regions of the home and
foreign country. In each region consumers obtain utility from non-traded goods and a traded good
aggregate.

In order to highlight the role of exchange rates in explaining the consumption-real exchange rate
correlation, we impose a lot of symmetry between regions and countries. Specifically, we assume
that factor markets are segmented both across regions and across countries; financial markets
are incomplete both within and across countries; and fiscal policies and productivity processes
are symmetric within and across countries. Clearly these assumptions limit risk-sharing within
countries and thus make it harder for us to establish the hypothesis that exchange rate movements
represent a major explanation of the consumption-real exchange rate anomaly. In the robustness
section we explore the role of these model features for the numerical results. For instance, we will
allow for productivity shocks to be more correlated within countries than across countries; and for
a possibility of fiscal transfers across regions within a country, but not across countries.\footnote{Since our focus is on business cycle dynamics of the model we limit our robustness analysis to productivity and fiscal transfers. Changes in labor mobility and financial markets incompleteness are typically long-term, structural phenomena that do not fluctuate at the business cycles frequency.}

It is convenient to set out the notation from the point of view of either region in the home
country. Let the utility of a representative infinitely-lived home household in region $j$ evaluated
from date 0 be defined as:

$$U_{0,j} = E_0 \sum_{t=0}^{\infty} \beta^t \left( \frac{C_{t,j}^{1-\sigma}}{1-\sigma} - \frac{N_{t,j}^{1+\phi}}{1+\phi} \right), \quad \beta < 1, \quad j = E, W.$$ \hfill (3.8)

where $C_{t,j}$ is the composite home consumption bundle in region $j$, and we have labeled the regions
as $E$ (e.g. east) and $W$ (west), and $N_{t,j}$ is home region $j$ labour supply. Composite aggregators
are defined as:

\[ C_{t,j} = \left( \varphi \frac{1}{2} C_{Tt,j}^{1-\frac{1}{\varphi}} + (1 - \varphi) \frac{1}{2} C_{Nt,j}^{1-\frac{1}{\varphi}} \right)^{\frac{1}{\varphi-1}}, \quad C_{Tt,j} = \left( \frac{v}{2} \right)^{\frac{1}{\chi}} C_{Ht,j}^{1-\frac{1}{\chi}} + \left( 1 - \frac{v}{2} \right)^{\frac{1}{\chi}} C_{Ft,j}^{1-\frac{1}{\chi}} \right)^{\frac{1}{\chi-1}}, \]

where \( C_{Tt,j} \) and \( C_{Nt,j} \) represent respectively, the composite consumption of traded and non-traded goods. The elasticity of substitution between traded and non-traded goods is \( \varphi \), and the weight of tradable consumption in the aggregate consumption bundle is \( \varphi \). Traded goods consumption is decomposed into consumption of home and foreign good aggregates \( C_{Ht,j} \) and \( C_{Ft,j} \), where \( \frac{v}{2} \) captures a degree of home bias in the traded consumption aggregator, since \( v \geq 1 \). We further separate the consumption of home and foreign traded goods into consumption of region \( E \) and \( W \) goods, so for instance, home consumption of traded goods is defined as

\[ C_{Ht,j} = \left( \frac{1}{2} \right)^{\frac{1}{\chi}} \left( C_{HEt,j}^{1-\frac{1}{\chi}} + C_{HWt,j}^{1-\frac{1}{\chi}} \right)^{\frac{1}{\chi-1}}. \]

The foreign traded goods consumption is defined symmetrically. We abstract from a regional bias in consumption aggregators, since we do not have detailed data on regional consumption preferences.

The parameter \( \lambda \) is the elasticity of substitution between the home and foreign traded good. The elasticity of substitution across regional goods is captured by the parameter \( \chi \). In each case, we assume that consumption sub-aggregates are differentiated across the consumption of individual goods, with elasticity of substitution across varieties greater than one.

These consumption aggregates imply the following price index definitions:

\[ P_{t,j} = \left( \varphi P_{Tt,j}^{1-\varphi} + (1 - \varphi) P_{Nt,j}^{1-\varphi} \right)^{\frac{1}{1-\varphi}}, \]
\[ P_{Tt,j} = \left( \frac{v}{2} P_{Ht,j}^{1-\lambda} + \left( 1 - \frac{v}{2} \right) P_{Ft,j}^{1-\lambda} \right)^{\frac{1}{1-\lambda}}, \]
\[ P_{Ht,j} = \left( \frac{1}{2} \right)^{\frac{1}{1-\chi}} \left( P_{HEt,j}^{1-\chi} + P_{HWt,j}^{1-\chi} \right)^{\frac{1}{1-\chi}}, \quad j = E, W. \]
where \( P_{t,j} \) is the price of the home aggregate consumption bundle; \( P_{Tt,j} \) and \( P_{Nt,j} \) represent traded and non-traded region \( j \) price levels; and \( P_{Ht,j} \) and \( P_{Ft,j} \) are retail prices of region \( j \) home and foreign traded goods, respectively. In turn, \( P_{Ht,j} \) is broken down into prices of region \( E \) and region \( W \) home traded good.

We define the intra-national real exchange rate in each country as the ratio of prices of regional consumption bundles. Specifically, in country \( H \), the intra-national real exchange rate is \( P_{Wt}/P_{Et} \), while in country \( F \), it is \( P_{Wt}^* / P_{Et}^* \). The cross country real exchange rate for region \( j \) is defined as the relative price of the foreign consumption bundle to the home bundle for that region:

\[
RER_{t,j} = \frac{S_t P_{t,j}^*}{P_{t,j}}.
\]

Here \( S_t \) is the nominal exchange rate, and \( P_{t,j}^* \) is the price of foreign region \( j \) consumption bundle.

In principal, there are four different cross country real exchange rates. However, as we describe below, the model dichotomizes so that within country and across country variables can be solved separately. Hence we can describe a single cross country real exchange rate as an aggregator of the above expressions.

We assume that each region within each country has access to two non-contingent one-period nominal bonds – a domestic currency denominated bond and a foreign currency denominated bond.

Thus, financial markets are incomplete both within countries and across countries.\(^ {34} \)

\(^{34}\)Since the number of shocks in the model far exceeds the number of assets available, the spanning condition is not satisfied, and financial markets are incomplete in our model. The international asset structure assumed in this paper is quite standard in the New Open Economy Macroeconomics literature (see Kollmann (2002), Benigno and Thoenissen (2008), and many others). However, the interaction of the asset market structure with monetary policy specification and their joint effects on the consumption-real exchange correlations are investigated less. Benigno and Küçük (2012), for instance, look at the model with two international currency bonds and show that the monetary policy specification affects the degree of internation risk sharing. Our work adds to this literature by focusing on the implications of the exchange rate regime for the consumption-real exchange rate correlation.
Fiscal authorities within countries can engage in some income redistribution. The expression $\Gamma_{t,j}$ represents net transfers to region $j$ households from the home government. A large literature in public finance has measured the extent of fiscal transfers across states or provinces within countries in response to regional income differentials (see e.g. [Boadway and Shah (2009)]). Net transfers may take place through endogenous variations in regional tax payments or explicit transfers in response to temporary or permanent shocks to relative regional GDP. Conceptually, such transfers should be distinguished from regional government spending initiatives, which may also contain a macroeconomic stabilization objective. The nature and calibration of the within country fiscal transfers are discussed in more detail below. We set $\Gamma_{t,j}$ to 0 in the benchmark calibration, but in the robustness section, we allow for a degree of within-country risk sharing generated by these fiscal transfers.

Thus, the home region $j$ budget constraint is given by:

$$P_{t,j}C_{t,j} + \frac{B_{t+1,j}}{\Theta_{j,H}(\frac{B_{t+1,j}}{P_{t,j}})} + \frac{S_{t}B_{t+1,j}^*}{\Theta_{j,F}(\frac{S_{t}B_{t+1,j}^*}{P_{t,j}})} = W_{t,j}N_{t,j} + \Pi_{t,j} + R_{t}B_{t,j} + R_{t}^{*}S_{t}B_{t,j}^* + \Gamma_{t,j}, \quad j = E, W. \quad (3.9)$$

where $B_{t+1,j}$ ($B_{t+1,j}^*$) indicates home (foreign) currency bond holdings of the home household in region $j$, $R_{t}$ ($R_{t}^*$) is the nominal interest rate on home (foreign) currency-denominated bonds, $W_{t,j}$ represents the nominal wage, assumed to be equalized across sectors within regions, and $\Pi_{t,j}$ represents profits from the home regional firm.

The functions $\Theta_{j,H}(B_{t,j}/P_{t,j})$ and $\Theta_{j,F}(S_{t}B_{t+1,j}^*/P_{t,j}^*)$ which satisfy $\Theta_{j,H}(\bar{B}/\bar{P}) = 1$, $\Theta'_{j,H} > 0$, $\Theta''_{j,H} > 0$ and $\Theta_{j,F}(\bar{S}\bar{B}^*/\bar{P}^*) = 1$, $\Theta'_{j,F} > 0$, $\Theta''_{j,F} > 0$, represent the presence of transactions costs for bond holdings, where we assume there are separate cost functions for the holdings of domestic and foreign bonds, with $\bar{B}/\bar{P}$ denoting a steady state real value of domestic bond holdings, and analogously for steady state foreign bond holdings. The assumptions of a cost of foreign bond
holdings is standard approach to induce stationarity in open economy models with incomplete markets (see Schmitt-Grohe and Uribe (2003)).

Households choose consumption of individual goods, labor supply in each sector and region, and bond holdings in the usual way. The Euler equation for nominal bond pricing in region \( j \) of the home country is given by:

\[
\frac{C_{t,j}^{\sigma}}{P_{t,j}} \Theta_{j,H}(\cdot) = R_{t+1} E_t \beta \frac{C_{t+1,j}^{\sigma}}{P_{t+1,j}}, \quad j = E, W. \tag{3.10}
\]

The optimal consumption-leisure trade-off is described by:

\[
W_{t,j} = C_{t,j}^{\sigma} N_{t,j}^{\phi} P_{t,j}, \quad j = E, W. \tag{3.11}
\]

Foreign household’s preferences and choices can be defined exactly symmetrically.

The intra-national risk-sharing condition for regions \( E \) and \( W \) within the home country is

\[
E_t \frac{C_{t+1,E}^{\sigma}}{C_{t,E}^{\sigma}} \frac{P_{t,E}}{P_{t+1,E}} \Theta_{E,H}(\cdot) = E_t \frac{C_{t+1,W}^{\sigma}}{C_{t,W}^{\sigma}} \frac{P_{t,W}}{P_{t+1,W}} \Theta_{W,H}(\cdot) \tag{3.12}
\]

The international risk-sharing condition for region \( j \) in the home country and region \( i \) in the foreign country is

\[
E_t \frac{C_{t+1,j}^{\sigma}}{C_{t,j}^{\sigma}} \frac{S_{t+1}}{S_t} \frac{P_{t,j}}{P_{t+1,j}} \Theta_{j,F}(\cdot) = E_t \frac{C_{t+1,i}^{\sigma}}{C_{t,i}^{\sigma}} \frac{P_{t,i}}{P_{t+1,i}} \Theta_{i,F}(\cdot) \quad i, j = E, W. \tag{3.13}
\]

This implies that up to a first-order approximation, expected consumption growth differentials across regions and countries are positively related to expected changes in the real exchange rate.

\[\text{35 With four separate consumer households (i.e. two countries, two regions), it is necessary to impose bond holding costs on only three of these in order to induce stationarity. Accordingly, we assume that both home regions and the foreign east region are subject to these costs.}\]
Firms in each sector and region produce using capital, $K$, and labor, $N$. A typical firm in the non-traded (traded) sector in region $j$ has production function $Y_{Nt,j}(i) = A_{Nt,j} N_{Nt,j}(i)^{(1-\alpha)} K_{Nt,j}(i)^{\alpha}$, $(Y_{Ht,j}(i) = A_{Ht,j} N_{Ht,j}(i)^{(1-\alpha)} K_{Ht,j}(i)^{\alpha})$. Thus, there are two technology shocks in each region – shocks to the non-traded sector $A_{Nt,j}$, and to the traded sector $A_{Ht,j}$. These shocks play substantially different roles in determining the consumption-real exchange rate correlation.

Firms choose their path of investment, $I$, so as to attain an optimal capital stock, where the relationship between investment and capital for a non-traded (traded) firm is $K_{Nt+1,j}(i) = (1 - \delta_K) K_{Nt,j}(i) + \phi(I_{Nt,j}(i)/K_{Nt,j}(i)) K_{Nt,j}(i)$ $(K_{Ht+1,j}(i) = (1 - \delta_K) K_{Ht,j}(i) + \phi(I_{Ht,j}(i)/K_{Ht,j}(i)) K_{Ht,j}(i))$. Here $\phi(.)$, satisfying $\phi'(.) > 0, \phi''(.) < 0$, represents an adjustment cost technology for investment incurred directly by the firm. Finally, $\delta_K$ is the depreciation rate for capital.

We assume that prices in each sector and region are sticky, and firms re-set prices according to a Calvo pricing policy, where the probability of re-adjusting the price is sector specific, and is denoted as $1 - \kappa_k$ for sector $k$ in each period (the parameters are identical across regions). Newly price-setting firms follow one of two policies. Some firms reset prices in the standard forward looking manner, but a fraction of firms re-set prices according to a ‘rule of thumb pricing’ as in [Gali and Gertler (1999)]. This has the effect of making the dynamics of sectoral inflation partly backward looking. This generalized specification for price setting allows us to address a well known critique of the Calvo model that it produces too much variability in inflation in response to new information - see Gali and Gertler (1999), Gali, Gertler, and Lopez-Salido (2001), McAdam and Willman (2004), and Christiano, Eichenbaum, and Evans (2005). As we shall show below, this will
also be important in accounting for the impact of the nominal exchange rate on the cross-country consumption correlation.

For forward looking price setting firms, the optimal price for a typical firm $i$ in sector $k$ of the home country region $j$ is written as:

$$P_{k,t,j}^f(i) = \frac{E_t \sum_{z=0}^{\infty} m_{t+z,j} \kappa_k^z MC_{k,t+z,j} Y_{k,t+z,j}(i)}{E_t \sum_{z=0}^{\infty} m_{t+z,j} \kappa_k^z Y_{k,t+z,j}(i)},$$

(3.14)

where $m$ is a stochastic discount factor defined as $m_{t+z,j} = \beta^z \left(\frac{C_{t+z,j}^{-\sigma}}{C_{t,j}^{-\sigma}}\right) \left(P_t/P_{t+1}\right)$, $Y_{k,t,j}(i)$ is demand for firm $i$'s good in sector $k$, region $j$, and $MC_{k,t,j}$ is the nominal marginal cost.\textsuperscript{38}

For firms that follow a rule of thumb pricing policy, we follow Gali and Gertler (1999) in assuming that they set the price equal to last period’s newly set price adjusted for last period’s inflation rate. Thus, the rule of thumb adjusted price is

$$P_{k,t,j}^b = \bar{P}_{k,t-1,j} \frac{P_{t-1,j}}{P_{t-2,j}},$$

(3.15)

where $\bar{P}_{k,t}$ represents the average newly set price in sector $k$. If the fraction of rule of thumb firms is given by $\omega$, then the newly set price in sector $k$ is

$$\bar{P}_{k,t,j} = \left[(1 - \omega) P_{k,t,j}^f(1-\theta) + \omega P_{k,t,j}^b(1-\theta)\right]^{\frac{1}{1-\theta}},$$

(3.16)

where $\theta > 1$ is the elasticity of substitution across varieties.

In the aggregate, the price index for the home good in each sector, in each region then follows include the sticky information model of Mankiw and Reis (2002) and the model of Christiano, Eichenbaum, and Evans (2005), which assumes both backward indexation in price adjustment, as well as the restriction that prices must be set before current information on shocks is revealed. Our results remain robust to these approaches.\textsuperscript{38} We assume that there is an optimal subsidy in place that eliminates the distortionary effect of the price markup in each sector. Note that in equilibrium all firms within a sector will adjust to the same price.
the process given by:

\[
P_{k,t,j} = [(1 - \kappa)P_{k,t,j}^{1-\theta} + \kappa P_{k,t-1,j}^{1-\theta}]^{1-\varpi}. \tag{3.17}
\]

The behavior of foreign firms and the foreign good price index may be described analogously.

We assume that the home country monetary authority follows a Taylor rule, which targets both the CPI inflation rate \(\Pi_t = P_t/P_{t-1}\) and the output gap \(\tilde{Y}_t\). Here the inflation rate is an equally weighted geometric average of the inflation rate in regions \(E\) and \(W\) in the home country, and the output gap is also an equally weighted average of GDP less the flexible price GDP in each region. In addition, the monetary rule puts some weight on the nominal exchange rate between the home and foreign country. Specifically, the nominal interest rate \(R_{t+1}\) is set so that:

\[
R_{t+1} = \beta^{-1} (\Pi_t)^{\gamma_{\pi}} (\Pi_{St})^{\delta} \left(\tilde{Y}_t\right)^{\gamma_y} \tag{3.18}
\]

where \(\Pi_{St} = S_t/S_{t-1}\) is the gross rate of change of the nominal exchange rate. \(\gamma_{\pi}\) is the weight that monetary authority places on domestic inflation, \(\delta\) is the weight on exchange rate growth rate, and \(\gamma_y\) is the weight placed on the output gap.\(^{39}\) For very high values of \(\delta > 0\), the exchange rate will be effectively fixed.\(^{40}\) The main issue of interest here will be to contrast the behavior of consumption and real exchange rates across countries for values of \(\delta = 0\), which we associate with a regime of inflation targeting in both countries, with a situation where the value of \(\delta\) goes to a very high positive number, which we associated with a fixed exchange rate regime.

\(^{39}\)Our specification for the monetary policy rule is motivated by its empirical relevance (see Taylor (1993), Taylor (2001), Clarida, Gali, and Gertler (2000) and many others). Several papers have considered the theoretical implications of alternative specifications of the policy rules for the dynamics of the real exchange rate (see, for instance, Benigno (2004), Benigno and Benigno (2008), Benigno and Küçük (2012). For instance, Benigno and Küçük (2012) study a policy rule that targets tradable good price inflation and show that in this case the nominal exchange rate is primarily determined by the terms of trade, as opposed to the real exchange rate. The real exchange rate becomes more relevant under the monetary policy rule that targets CPI inflation. In such a scenario the predicted risk-sharing will also be higher. Our primary interest is in contrasting the effects of fixed versus flexible exchange rate regime under the same monetary policy rule for price inflation. We have also considered the case with interest rate inertia in the Taylor rule but found that results remained largely unchanged.

\(^{40}\)Note that this specification is not associated with the problems of indeterminacy of a fixed exchange rate regime. The nominal exchange rate is fully determined, given an initial value \(S_0\). See Benigno, Benigno, and Ghironi (2007).
The foreign monetary authority follows an identical Taylor rule, except it does not directly target the exchange rate.\footnote{In fact, for the value of \(\text{corr}(\Delta c - \Delta c^*, \Delta RER)\), it is irrelevant which monetary authority targets the exchange rate, or whether it is done jointly. It is convenient to focus solely on home country exchange rate targeting, because this makes the analysis of nominal price adjustment easier to exposit.}

Market clearing conditions are defined in the standard way. Goods market clearing conditions for region \(j\) are given as:

\[
Y_{Ht,j} = C_{Ht,E} + C_{Ht,W} + C_{Ht,E}^* + C_{Ht,W}^* + I_{Ht,E} + I_{Nt,E},
\]

\[
Y_{Nt,j} = C_{Nt,j}, \quad j = E, W.
\]

3.2 The consumption-real exchange rate correlation in the model

Below we will use lowercase letters to denote log transformations of all variables. The (log) linear approximation to equation \((3.12)\) is given by:

\[
E_t(\Delta c_{t+1,E} - \Delta c_{t+1,W}) = \frac{1}{\sigma} E_t \Delta \text{rer}_{t+1,WE} + \frac{1}{\sigma} \bar{\sigma} b_{t+1,E}
\]

where \(\Delta c_{t+1,j} = c_{t+1,j} - c_{t,j}, \quad j = E, W; \Delta \text{rer}_{t+1,WE} = \text{rer}_{t+1,WE} - \text{rer}_{t,WE};\) and \(\bar{\sigma}\) is a constant coefficient. This captures the within-country relationship between expected consumption growth and expected real exchange rates movements, where \(\text{rer}_{t,WE}\) is the real exchange rate between west and east regions in the home country, i.e. \(\text{rer}_{t,WE} = p_{t,W} - p_{t,E}\). The last term on the right hand side represents the wedge in consumption growth, for given real exchange rates, associated with differential home country costs of foreign bond holdings. This represents a very small term, and has no first-order implication for the relationship between relative consumption and real exchange rates.
The analogous expression between countries from (3.13) is

\[ E_t(\Delta c_{t+1,j} - \Delta c_{t+1,i}^*) = \frac{1}{\sigma} E_t \Delta rer_{t+1,j,i}^* + \frac{1}{\sigma} \Theta_{t+1,j}^* b_{t+1,j}^* , \quad j, i = E, W \]  \hspace{1cm} (3.22)

where \( rer_{t+1,j,i}^* \) is the real exchange rate between region \( j \) in the home country and region \( i \) in the foreign country.

In this model, a negative correlation between relative consumption and the real exchange rate can occur, even though in expectations, there is a positive relationship between the two variables. To show this, let’s focus, for instance, on equation (3.22) and define \( y_{t+1} \equiv \Delta c_{t+1,j} - \Delta c_{t+1,i}^* \), and \( x_{t+1} \equiv (1/\sigma) \Delta rer_{t+1,j,i}^* \). Then, ignoring the bond market adjustment term, the risk sharing condition can be stated as

\[ y_{t+1} = x_{t+1} + \nu_{t+1} \]  \hspace{1cm} (3.23)

where \( \nu_{t+1} = y_{t+1} - E_t y_{t+1} - (x_{t+1} - E_t x_{t+1}) \) and the regression coefficient implied by (3.23) is

\[ \hat{\beta}_{x,y} = 1 + \frac{\text{Cov}(x_{t+1}, \nu_{t+1})}{\text{Var}(x_{t+1})} \]

This coefficient can be positive or negative depending on the covariance between the innovations to relative consumption growth less real exchange rate growth, \( \nu_{t+1} \), and the real exchange rate. Focusing on cross-country relationship between consumption and real exchange rate, Benigno and Thoenissen (2008) showed that productivity shocks to the traded goods sector, in an environment of incomplete financial markets, can generate a negative correlation between the two variables. Given our empirical findings of negative correlation across countries (that also depends on the exchange rate regime), but positive correlation within countries, the challenge thus, lies in accounting for all these facts simultaneously. We argue that the key is in the behavior of various components of the
real exchange rates.

The within-country real exchange rate in the model can be decomposed into a function of two relative prices:

$$rer_{t,WE} = \rho(\tau_{Nt,W} - \tau_{Nt,E}) + \rho \tau_{t,WE}$$  (3.24)

Here $\tau_{Nt,W} = p_{Nt,W} - p_{Ht,W}$, $\tau_{Nt,E} = p_{Nt,E} - p_{Ht,E}$, and $\tau_{t,WE}$ is the within country terms of trade, defined as $p_{Ht,W} - p_{Ht,E}$. The within-country real exchange rate then moves for two reasons. First, it is affected by regional differences in internal relative prices. If the west region relative price of non-traded goods rises, relative to the east region relative price, there will be an east region real depreciation. Secondly, the real exchange rate is affected by the within-country terms of trade $\tau_{t,WE}$. For a given value of $\tau_{Nt,W} - \tau_{Nt,E}$, a terms of trade deterioration causes an east region real exchange rate depreciation, since it raises the west relative to the east price of consumption. It is the first term, capturing the response of relative prices of non-traded goods, that allows for relative consumption and real exchange rates to co-vary negatively in the theoretical model.

International real exchange rate can be decomposed into a function of three relative prices:

$$rer_{t} = \rho(\tau_{Nt}^{*} - \tau_{Nt}) + (1 - \rho)(v - 1)\tau_{t} + \rho \tau_{t}$$  (3.25)

where in each case, the relative prices are defined as the country averages across regions. Hence, $\tau_{Nt} = p_{Nt} - p_{Ht}$, $\tau_{Nt}^{*} = p_{Nt}^{*} - p_{Ht}^{*}$ and $\tau_{t}$ is the international terms of trade, defined as $s_{t} + p_{Ft}^{*} - p_{Ht}^{*}$.

The interpretation for equation (3.25) is the same as for (3.24), except for the presence of the additional term linking the terms of trade with the real exchange rate, due to home country bias $(v \geq 1)$ in the consumption basket for each country.

42To be more precise, in the cross-country case, the definition of the real exchange rate depends on the comparison between regions located in the two countries. To simplify the reporting of results we define the cross country real exchange rate as the regional average compared across countries.
To see how technology shocks affect relative consumption, the real exchange rate and its components, we next calibrate and simulate the model. We consider the world economy as consisting of two symmetric countries, and within each country there are two equally sized regions, matching the properties of the US economy in quarterly data. Most of the preference parameter values are standard in the literature and, in particular, follow closely those adopted by Stockman and Tesar (1995). Parameter values used in the calibration are summarized in Table 9.

Table 9: Benchmark model parameters

<table>
<thead>
<tr>
<th>parameter</th>
<th>value</th>
</tr>
</thead>
<tbody>
<tr>
<td>subjective discount factor</td>
<td>$\beta$</td>
</tr>
<tr>
<td>labor supply elasticity</td>
<td>$\phi$</td>
</tr>
<tr>
<td>risk aversion</td>
<td>$\sigma$</td>
</tr>
<tr>
<td>elasticity of substitution across regional traded goods</td>
<td>$\chi$</td>
</tr>
<tr>
<td>elasticity of substitution b/n N and T goods</td>
<td>$\varphi$</td>
</tr>
<tr>
<td>elasticity of substitution b/n H and F goods</td>
<td>$\lambda$</td>
</tr>
<tr>
<td>T goods weight in aggregate consumption</td>
<td>$\eta$</td>
</tr>
<tr>
<td>home bias in T goods sector</td>
<td>$\nu$</td>
</tr>
<tr>
<td>depreciation rate</td>
<td>$\delta_K$</td>
</tr>
<tr>
<td>capital share</td>
<td>$\alpha$</td>
</tr>
<tr>
<td>1-probability of re-adjusting price in sector N</td>
<td>$\kappa_N$</td>
</tr>
<tr>
<td>1-probability of re-adjusting price in sector T</td>
<td>$\kappa_T$</td>
</tr>
<tr>
<td>weight on inflation in interest rate rule</td>
<td>$\gamma_i$</td>
</tr>
<tr>
<td>weight on the output gap in the interest rate rule</td>
<td>$\gamma_y$</td>
</tr>
<tr>
<td>weight on exchange rate growth in interest rate rule</td>
<td>$\delta$</td>
</tr>
<tr>
<td>fraction of rule of thumb firms</td>
<td>$\omega$</td>
</tr>
<tr>
<td>investment adjustment costs</td>
<td>$\zeta_{ik}$</td>
</tr>
<tr>
<td>bond adjustment costs</td>
<td>$\zeta$</td>
</tr>
</tbody>
</table>

The discount factor is set at 0.99. The labor supply elasticity $\phi$ is set at unity. The elasticity of intertemporal substitution is set at 0.5 so that $\sigma = 2$. The elasticity of substitution between non-traded and traded goods is usually estimated to be substantially below unity, so we set $\varphi = 0.5$, midway between the estimates of Stockman and Tesar (1995) and Mendoza (1995). The elasticity of substitution between home and foreign goods is set at $\lambda = 2.5$, a benchmark estimate. In the analysis below we also show how the results are affected by variations in $\lambda$. In addition, without a more precise measure, we assume that the elasticity of substitution across regional traded goods within each country is set at the same value, so that $\chi = \lambda$. We assume that the non-traded goods

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$^{43}$See Ruhl (2005), Matsumoto (2007) among others.
sector is 60% of GDP, somewhat smaller than that of the US economy, but larger than that for smaller economies. This pins down parameter $\varrho$. The degree of home bias in the traded goods sector is set at $v = 1.25$, which in steady state implies that imports are 15% of GDP. We follow Ghironi and Melitz (2005) in setting the elasticity of the bond adjustment cost equal to 0.0025. The elasticity of the marginal investment adjustment cost function at the steady state is set to reproduce a ratio of investment to output volatility approximately equal to 3. In the baseline setting, this implies that $(\bar{\sigma'}/i/k) / (\bar{\sigma'}) = -0.75$. The depreciation rate on capital is set to equal 0.025 at quarterly frequency. Capital share in the production function is set to 0.33.

The degree of price rigidity tends to be substantially higher in the non-traded goods sector than in the traded goods sector. Nakamura and Steinsson (2008) measure the median duration of fixed prices in the US service sector to be 5 quarters, and 3 quarters for the non-service sector. We use these measures to set $k_N = 0.8$ and $k_T = 0.67$. The size of the ‘rule of thumb’ sector in price setting determines the degree to which there is a backward looking coefficient in the inflation equations. We choose $\omega$ to match the estimates of Gali and Gertler (1999), who find that the backward looking coefficient is around 0.3. This implies that just over a quarter of firms in each sector use a rule of thumb approach to price setting rather than a forward looking approach. With this value of $\omega$ our model also matches the hybrid Phillips curve as in Gali and Gertler (1999) (it contains both forward and backward components of inflation).

We set the coefficients on the Taylor rule for monetary policy to be standard values of $\gamma_\pi = 1.5$ and $\gamma_y = 0.25$. Under a fixed exchange rate, $\delta$ is set at a high value so that the nominal exchange

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44 There is a wide variety of estimates of the size of the backward looking component in inflation determination. Gali, Gertler, and David Lopez-Salido (2005) find that, in a model where the output gap is proxied by labour share, the New Keynesian Phillips curve is estimated to be driven mostly by forward looking inflation expectations. But their estimation approach has been questioned by Rudd and Whelan (2006), who argue that model misspecification and omitted variable bias imply that these estimates are consistent with a large role for backward looking price setting. Linde (2005), using a full information maximum likelihood approach, estimates the shares of backward and forward looking pricing to be roughly equal. In what follows we assess the robustness of our results with respect to this parameter.
rate is constant (the actual value for $\delta$ is irrelevant once it is set high enough).

In measuring the persistence and volatility of productivity shocks to the traded and non-traded sector in each region, we follow Benigno and Thoenissen (2008). They estimate home traded sector productivity shocks to have persistence $\mu_H = 0.85$ and standard deviation $\sigma_H = 1.94\%$, while non-traded shocks are much less persistent and less volatile, with $\mu_N = 0.3$ and $\sigma_N = 0.71$. The foreign regional productivity process is symmetric. In addition, we impose the covariance structure of shocks as estimated by Benigno and Thoenissen (2008)\footnote{In particular, taking the shock innovations at $\varepsilon = [\varepsilon_H, \varepsilon_F, \varepsilon_N, \varepsilon_N]',$ we use $\text{Var}(\varepsilon) = \begin{bmatrix} 3.76 & 1.59 & 0.72 & 0.44 \\ 1.59 & 3.76 & 0.44 & 0.72 \\ 0.72 & 0.44 & 0.51 & 0.21 \\ 0.44 & 0.72 & 0.21 & 0.51 \end{bmatrix}$.} In our benchmark calibration of productivity we assume that regional productivities in traded and non-traded sectors within a country are characterized by the same shock processes as the same productivities across countries. But it is likely that there is higher shock correlation within countries than across countries. Kollmann (1995), finds that the correlation of productivity shocks across US regions is about 50% higher than that across G7 economies. Accordingly, we also consider a robustness parameterization for productivity shocks in which the within-country correlation of tradable and non-tradable productivity shocks is increased by 50%.

As for the fiscal policy, as discussed above, we assume that in the benchmark calibration, the fiscal policy is symmetric within and across countries. This implies that the transfer function is $\Gamma_{t,j} = 0$. In the robustness exercise we allow for within-country fiscal risk sharing. In calibrating transfer function $\Gamma_{t,j}$ in this case, we rely on available estimates in the data. For instance, Hagen (2007) and Poghosyan, Senhadji, and Cottarelli (2015) summarize the empirical literature on the size of interregional transfers across OECD countries. Sachs and Sala-i Martin (1992) provide early estimates for the US, and suggest that a fall in regional income (using 9 US regions) of one dollar

\footnote{Note that the consumption real exchange rate correlation depends only on relative shocks. Hence what matters for measuring the correlation is the relative size of traded to non-traded volatility and the covariance pattern, and not the absolute scale of volatility.}

\[42\]
elicits a net transfer of between 33 and 40 cents towards the region. Subsequently however the literature distinguished between fiscal redistribution, designed to correct long run differences in GDP per capita across regions, and fiscal risk-sharing, which uses taxes or transfers to correct cyclical differentials in income across regions. It is the latter concept that is relevant for our model. Estimates of fiscal risk-sharing vary across the literature, ranging from 31% in Canada (see Bayoumi and Masson (1995)) to 3% in Italy (Obstfeld and Peri (1998)). Estimates for the US typically range from 10 to 20%, although using a Pooled Mean Group estimator, Poghosyan, Senhadji, and Cottarelli (2015) estimate a risk sharing coefficient of 0.11 for the US. In our model, the risk sharing coefficient is measured by the parameter obtained from a log linear approximation of the $\Gamma_t = \Gamma(Y_{i,t}/Y_{j,t})$ and Sachs and Sala-i Martin (1992) suggest that a rise in regional income of one dollar increases transfers out of the region by 20 cents, and symmetrically for a fall in regional income. Other estimates broadly support the existence of endogenous fiscal transfers of this magnitude. Extrapolating from these empirical estimates, this implies an elasticity of $\Gamma_{t,j}$ to relative income between regions equal to 0.2 in our robustness exercise.

### 3.3 Within and Across country Dichotomy

Given the symmetry of the model, it is possible to separately characterize the impulse responses for within-country and across country variables. As shown in the appendix, we can solve for relative consumption responses and real exchange rates between region $E$ and $W$ in each country (as well as relative responses for all other variables) as functions of relative productivity shocks between $E$ and $W$, without information on the absolute size of individual $E$ and $W$ productivity shocks, or productivity shocks in the partner country. In addition, because regions within a country obviously share the same monetary policy, within-country relative prices and quantities do not depend on the exchange rate regime or the monetary rule. Likewise, the model allows for aggregation so that we
may solve for the responses of relative consumption, real exchange rates and relative magnitudes of all other variables compared across countries as functions of relative productivity shocks in the home and foreign country, without knowledge of the regional distribution of the shocks. Hence, the model can be simulated separately for within and across country moments, using the calibrated parameter assumptions and the different variance covariance matrices of productivity shocks.

3.4 Benchmark results

Since the properties of the model have been extensively evaluated in the international business cycle literature (see, for instance, Stockman and Tesar (1995), Benigno and Thoenissen (2008)), we focus only on the model predictions for the risk-sharing measures relevant to our empirical analysis. Figure 1 presents the impulse responses to a negative shock to traded sector productivity in the home country.

Panel (a) of Figure 1 illustrates the responses of home relative consumption and real exchange rate. The decomposition of real exchange rate adjustment is illustrated in Panel (b) of the Figure. It shows the response of $\tau_N - \tau_N$, the difference between internal relative prices in the two countries, and the response of $\tau$, the home terms of trade (as shown in equation (3.25)). Panel (c) shows the alternative decomposition of the real exchange rate, breaking down the response between the nominal exchange rate growth, $\Delta s$, and the response of relative CPI inflation, $\Delta p^* - \Delta p$. Finally, panel (d) shows the response of domestic inflation in the home traded good $\Delta p_H$ and the home non-traded good $\Delta p_N$. Both figures contrast the dynamics under a Taylor rule (‘flex’) with those under an exchange rate peg (‘fixd’).

Panel (a) of Figure 1 establishes that in both cases – the Taylor rule policy and fixed exchange rate – the decline in home traded goods productivity leads to a fall in relative consumption, and a
Notes: The figures present the impulse responses of the home country's relative consumption, international RER and its components to a negative tradable sector productivity shocks under benchmark model calibration. The results are presented under both flexible (flex) and fixed (fixd) exchange rate regimes.

real exchange rate depreciation. The real depreciation is achieved by a rise in $\tau_N^* - \tau_N$, the ‘relative-relative’ price of non-traded goods across countries. The income effect of a fall in productivity leads to a fall in demand for non-traded goods in the home country, and thus a fall in $\tau_N$, implying a real exchange rate depreciation. At the same time, there is a fall in $\tau$, i.e. a terms of trade appreciation, as the global supply of the home traded good is reduced. This generates a real exchange rate appreciation. Panel (a) shows that the rise in $\tau_N^* - \tau_N$ dominates the fall in $\tau$, so the net effect amounts to a real exchange rate depreciation.

The responses to the shock differ across exchange rate regimes. Panel (a) indicates that, relative to the exchange rate peg, relative consumption falls by less, and the real exchange rate depreciates by more under a Taylor rule. This greater depreciation is achieved by a bigger increase in $\tau_N^* - \tau_N$,  

and a smaller terms of trade appreciation under a Taylor rule. The key reason for the difference in responses is the ability of the nominal exchange rate to respond immediately with the Taylor rule. Panel (c) clearly illustrates this difference. With the Taylor rule, there is a nominal exchange rate depreciation. While there is a slight fall in $\Delta p^* - \Delta p$, this is fully offset by the nominal exchange rate response, facilitating a real depreciation. Under an exchange rate peg, however, all the real depreciation must be achieved by relative domestic deflation, or a rise in $\Delta p^* - \Delta p$. The forces for deflation can be seen in panel (d), indicating that the fall in traded sector productivity leads to a fall in non-traded goods inflation, and a slight rise in home traded goods inflation. Under the exchange rate peg, there must be greater deflation than under a Taylor rule, in order to sustain a real exchange rate depreciation.

The upshot of this discussion is that, while shocks to productivity in the traded sector can lead to a negative co-movement between relative consumption and the real exchange rate, this co-movement is not independent of the nominal exchange rate policy. Under a pegged exchange rate, the movement in the real exchange rate is limited, while under a Taylor rule, the immediate response of the nominal exchange rate facilitates a much larger immediate response in the real exchange rate.

Figure 2 gives the same breakdown for the case of a negative shock to non-traded sector productivity. This shock reduces home relative consumption and causes a real exchange rate appreciation. This is shown in panel (a) of the Figure. In panel (b), we see that the real appreciation is achieved by a combination of a fall in $\tau_N^* - \tau_N$, and a small rise in $\tau$. Intuitively, the fall in productivity in non-traded goods directly drives up their relative price, which implies a real exchange rate appreciation. Again, as in the previous case, the response of the nominal exchange rate is important in facilitating the real appreciation. With a flexible nominal exchange rate, most of the appreciation is achieved by a nominal appreciation, with a small increase in $\Delta p^* - \Delta p$. When the exchange
rate is fixed, all the appreciation must be generated by a fall in $\Delta p^* - \Delta p$, i.e. a rise in domestic relative to foreign inflation. Note however that although the decomposition of the real exchange rate response in this case does depend on the monetary rule, the co-movement between relative consumption and the real exchange rate is positive in either case. It is also noteworthy to see that while the real exchange rate response to a non-traded productivity shock is greater under a flexible exchange rate, the difference is less than the response to a traded goods shock. In the former case, a greater part of the response depends on changes in internal relative prices, and hence is less contingent on nominal exchange rate changes.

Figure 2: Impulse responses following a negative shock to N sector productivity

Notes: The figures present the impulse responses of the home country’s relative consumption, international RER and its components to a negative non-tradable sector productivity shocks under benchmark model calibration. The results are presented under both flexible (flex) and fixed (fixd) exchange rate regimes.
These impulse responses from the model suggest that, when the dominant shocks arise from the traded goods sector, then there may arise a negative correlation between relative consumption and real exchange rates. Moreover, the correlation may be affected by the policy rule. Under a fixed exchange rate regime, the response of the real exchange rate is dampened after both traded and non-traded sector productivity shocks, but more-so for traded sector shocks. This suggests that the negative correlation induced by the traded-sector shocks under the Taylor rule becomes weaker under a peg, and in fact can turn positive.

Next, we turn to intra-national consumption and real exchange rate dynamics. Under our benchmark calibration which features full symmetry between regions and countries, we find that the responses of relative consumption and real exchange rate to a within-country traded or non-traded goods relative productivity shock are identical to the responses of across country relative consumption and real exchange rate obtained under fixed exchange rates. Therefore, the within-country consumption-real exchange rate correlation and across-country consumption-real exchange rate correlation under the fixed exchange rate regime are exactly the same in our benchmark model specification.

We next assess the quantitative model predictions for the consumption-real exchange rate correlations within and between countries, and under different exchange rate regimes. Table 10 reports the simulation results for the model under benchmark calibration. The table summarizes both the across-country (Intl) and within-country (Intra) values for consumption real exchange rate correlation, the standard deviation of relative consumption, and the real exchange rate, as well as output, investment, and the nominal exchange rate, under the case of a Taylor rule, and under the

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47 This result is in line with the findings in Benigno and Thoenissen (2008).

48 Another way to evaluate the model is by testing the key implication of the model for risk-sharing, given by equation (4.13). We generalize this equation by dropping the bond market adjustment terms and allowing the coefficient of relative risk aversion to be country-specific. We then test this equation for 8 OECD countries using the Generalized Method of Moments. We fail to reject the international risk-sharing condition for all bilateral and joint tests. The detailed results can be found in online appendix.
exchange rate peg. The last row of the table also includes the within-country statistics obtained under the robustness calibration (Intra (robustness)), where we allow for endogenous fiscal transfers and higher co-movement between productivity shocks within countries.

Across countries, in comparison to the Taylor rule, the fixed exchange rate leads to a rise in the volatility of relative consumption, output, and investment, and a fall in the volatility of the real exchange rate. Under a Taylor rule, the value of corr($\Delta c - \Delta c^*$, $\Delta rer$) is $-0.08$, approximately that we found in the data for floating exchange rate pairs of regions. By contrast, under a fixed exchange rate, the correlation is $0.06$. Thus, we find that the value of corr($\Delta c - \Delta c^*$, $\Delta rer$), in the model does depend on the nominal monetary rule. The correlation is positive under fixed exchange rates, and negative under flexible exchange rates.

Within countries, the correlation is equal to $0.06$ – its cross-country counterpart under the fixed exchange rate regime in the benchmark correlation, but increases to $0.18$ when we allow for endogenous fiscal transfers and higher co-movement between productivity shocks within countries. It is notable, however, that even with these features, the within-country risk sharing is far from the complete markets implication of a unit correlation. Overall, these results line up very well with our empirical findings.

Table 10: Simulation results

<table>
<thead>
<tr>
<th>Scenario</th>
<th>std.dev.($\Delta c - \Delta c^*$)</th>
<th>std.dev.($\Delta rer$)</th>
<th>corr($\Delta c - \Delta c^*$, $\Delta rer$)</th>
<th>std.dev.($y$)</th>
<th>std.dev.($I$)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Intl flex, ($\delta = 0$)</td>
<td>1.00</td>
<td>0.8</td>
<td>-0.08</td>
<td>1.78</td>
<td>4.45</td>
</tr>
<tr>
<td>Intl fixed ($\lim \delta \to \infty$)</td>
<td>1.09</td>
<td>0.66</td>
<td>0.06</td>
<td>1.85</td>
<td>4.53</td>
</tr>
<tr>
<td>Intra (benchmark)</td>
<td>1.09</td>
<td>0.66</td>
<td>0.06</td>
<td>1.85</td>
<td>4.53</td>
</tr>
<tr>
<td>Intra (robustness)</td>
<td>0.96</td>
<td>0.59</td>
<td>0.18</td>
<td>1.79</td>
<td>4.43</td>
</tr>
</tbody>
</table>

3.5 Robustness

In this section we evaluate the robustness of our findings by considering the effects of (i) allowing for endogenous fiscal transfers within countries; (ii) higher within-country correlation of productivity
shocks; (iii) changes in the elasticity of substitution between traded goods; (iv) different degree to which prices are backward looking.

First, we introduce endogenous fiscal transfers into our benchmark calibration. Panel (a) of Figure 3 compares the responses of cross-region relative consumption and real exchange rate to a traded goods relative productivity shock to the across country impulse responses of the two variables under fixed exchange rates. In this scenario, the only structural difference between regions and countries is the presence of endogenous fiscal transfers. Without the endogenous transfers the two responses are identical, as argued above. However, when a fall in regional relative productivity and income precipitates transfers across regions, relative consumption falls by less within countries than the analogous measure does across countries, and the within country real exchange rate depreciates by less than the across country real exchange rate.

Figure 3: Intra-national vs international impulse responses to a negative shock to T sector productivity

Notes: The figures present the impulse responses of the home country’s international and intra-national relative consumption, as well as international and intra-national RERs to a negative tradable sector productivity shock under benchmark model calibration in panel (a) and negative non-tradable productivity shocks in panel (b). The results for international variables are presented under fixed exchange rate regime.

Panel (b) of Figure 3 illustrates the same comparison in response to a within-country deterior-
ration in non-traded goods productivity. As in the case of the traded goods shock, the response of relative consumption is muted due to the endogenous fiscal transfer between regions. Indeed these effects lead to a significant increase in the within-country consumption-real exchange rate correlation.

Second, we allow for a greater correlation of regional tradable and non-tradable productivity shocks within countries than across countries. The responses of relative consumption and real exchange rates are very similar to those under endogenous fiscal transfers, but are smaller quantitatively (not shown). Thus, the within country risk-sharing improves under more correlated shocks, but the improvement is quantitatively small.

A number of studies (see, for instance, Corsetti, Dedola, and Leduc (2008)) have shown that the elasticity of substitution between traded goods is a key parameter affecting the consumption-real exchange rate correlation in the class of models considered here. Therefore, we next investigate the sensitivity of our results to the trade elasticity. Figure 4 illustrates the impact of a traded sector productivity shock for a value of $\lambda = 5$, a higher elasticity than in the benchmark case. Now in response to the shock, there is a smaller terms of trade appreciation, due to a higher elasticity of substitution between traded goods. This leads to a greater real exchange rate depreciation under flexible exchange rates, with the nominal exchange rate playing a greater role. Then fixing the exchange rate has a greater impact on the real exchange rate, because it is harder to achieve the required depreciation with domestic disinflation.

Panel (a) of Figure 5 illustrates the results for the consumption-real exchange rate correlation under a fixed and flexible exchange rate regimes, as a function of the size of the trade elasticity $\lambda$. We also show the relationship between $\lambda$ and the within-country real exchange rate correlation in

\textit{Again we assume that the intra-country elasticity is the same as $\lambda$.}
Figure 4: Impulse responses following a negative shock to T sector productivity with $\lambda = 5$

Notes: The figures present the impulse responses of the home country’s relative consumption, RER, and its components to a negative traded-sector productivity shock under an alternative parameterization of parameter $\lambda = 5$. The results are presented under both flexible (flex) and fixed (fixd) exchange rate regimes.

the same figure, but it is easy to see that the within correlation coincides with the across correlation under the fixed exchange rate for all values of $\lambda$. With $\lambda$ close to unity, the correlation is equal to unity both across and within countries, independent of the nominal exchange rate regime, since we have effectively complete markets in that case (see Cole and Obstfeld (1991)). As $\lambda$ increases, we find that the correlation declines, both within and across countries. Moreover, the across country correlation is lower for a flexible exchange rate policy than under fixed exchange rates. For values of $\lambda$ in the range of 2 to 3, we see that the across country correlation reverses sign with the exchange rate regime. As $\lambda$ rises above 3, the correlation becomes negative under both monetary rules. But it is still the case that the fixed exchange rate rule generates a higher (more positive) correlation than the rule which allows nominal exchange rates to adjust.
Figure 5: Consumption-RER correlation: Sensitivity to trade elasticity

(a) benchmark calibration  

(b) with fiscal transfers and high prod. corr

Notes: The figure presents a plot of international and intra-national C-RER correlations in the model for various values of parameter $\lambda$. Cross-country correlations are presented under both inflation targeting (flex) and fixed (fixd) exchange rate regimes, while intra-national correlations are shown under the scenario with endogenous fiscal transfers, and with both transfers and higher within country productivity correlations.

Panel (b) of Figure 5 compares the cross-country correlations (under flexible and fixed exchange rates) to the intra-national correlation in the presence of endogenous fiscal transfers and with positive within-country productivity correlations. The figure makes it clear that the endogenous fiscal transfers significantly increase the within country risk-sharing across the business cycle for all values of elasticity $\lambda$. When a greater co-movement between productivity shocks is added to the model the consumption-real exchange rate correlation increases further, but the increase is not very big. These results suggest that endogenous fiscal transfers play a crucial role in intra-national risk-sharing at the business cycle frequency.

Lastly, we consider the role played by the assumption of backward looking prices, implied by the parameter $\omega$. Figure 6 repeats Figure 5 under the two scenarios: panel (a) assumes that $\omega = 0$, so that there are no rule of thumb price setters; while panel (b) assumes that $\omega = 0.5$, so that half of new price setters use a rule of thumb. With no backward pricing we see the negative relationship
between \( \text{corr}(\Delta c - \Delta c^*, \Delta rer) \) and \( \lambda \) still prevails. Moreover, the correlation is still more negative under a Taylor rule than under a fixed exchange rate. But in contrast to the baseline case, the correlation is significantly negative for all empirically reasonable values of \( \lambda \), for both the Taylor rule and the pegged exchange rate. Intuitively, in the absence of backward price setters, inflation rates are fully forward looking in the Calvo model. This makes the real exchange rate sufficiently flexible that even under a fixed exchange rate, the consumption-real exchange rate correlation is negative over a greater range of values of \( \lambda \) than in the baseline model.\(^{50}\)

Figure 6: Consumption-RER correlation and trade elasticity: Sensitivity to backward-looking price setting

Notes: The figure presents a plot of international and intra-national C-RER correlations in the model for various values of parameter \( \lambda \) under \( \omega = 0 \) and \( \omega = 0.5 \). Correlations are presented under both inflation targeting (flex) and fixed (fixd) exchange rate regimes.

Finally, panel (b) of Figure 6 indicates that as backward looking pricing becomes greater, the dichotomy between the fixed and flexible exchange rate for the consumption-real exchange rate correlation becomes much more pronounced. In this panel we set \( \omega = 0.5 \), so half of new price setters use a rule of thumb. In this case, we see the correlation is always positive under a fixed exchange rate, and for values of \( \lambda \geq 2.5 \) always negative under a flexible exchange rate. The

\(^{50}\)This intuition also explains why adding some inertia in real exchange rate through an interest rate smoothing term in the Taylor rule while setting \( \omega = 0 \) is not enough to reproduce our results.
within-country consumption-real exchange rate correlation remains positive for all values of \( \lambda \) in the considered range, independent of the degree of backward looking pricing in the model.\(^{51}\)

4 Conclusions

In this paper we have explored the nature of consumption risk-sharing and real exchange rate movements within and between countries. Within-country risk-sharing displays a significant positive correlation between relative consumption and real exchange rates, as standard risk-sharing theory predicts. Across countries, this correlation is negative, thereby revealing a ‘border effect’. This border effect is substantially accounted for by movements in the nominal exchange rate. In our sample, country pairs with low nominal exchange rate variance tend to have positive consumption-real exchange rate correlations. We show that our empirical findings remain valid through various robustness checks, both for country heterogeneity and habit persistence in consumption. We then showed that the role for the nominal exchange rate in consumption risk-sharing is consistent with the results of standard sticky price open economy models, in which real allocations are affected by the exchange rate regime, and the model is extended to a framework with a fraction of rule of thumb price setting.

Our paper omits two areas of interest for further research. One is the nature of the within-country consumption-real exchange rate correlation. While we find evidence of risk sharing, it is substantially less than implied by frictionless within-country financial markets. A second area of interest is in understanding the full role of the border in cross-country risk-sharing. While we find that the nominal exchange rate accounts for a sizeable fraction of the border effect, it does not account for it all. There still remain aspects of cross-country interrelationships which

\(^{51}\)It is important to note that our results on risk-sharing are not driven by significant departures from the Phillips curve as estimated in [Gali and Gertler (1999)]. In fact, our model implies the coefficient on lagged inflation in the hybrid Phillips curve that is comparable that in [Gali and Gertler (1999)].
lead the consumption-real exchange rate correlation to fall, independent of nominal exchange rate movement. Identifying these mechanisms is a question we leave to future research.
A Appendix

A.1 Data sources

Canada: For Canada our data includes series for consumption and prices for 10 provinces and 3 territories over 1981-2007 period. All provincial data is obtained from CANSIM. For prices we used CANSIM Table 326-0021 "Consumer Price Index, by province"; for consumption, we used "Personal expenditure on consumer goods and services, Chained (2002) dollars" series. We converted consumption into real per capita terms by dividing the original series by CPI and population.

Germany: Our data for Germany covers 16 bundeslander over 1995-2007 period. Consumption series are obtained from Regional Accounts database collected by Statistische Ämter des Bundes und der Länder. The series is called "Private consumption expenditure - price-settled, concatenated - for each inhabitants in Germany 1991 to 2006 after Federal states year index (2000 = 100)". Price data is from GENESIS-Tabelle: 61111-0010. The series name is "Verbraucherpreisindex für Deutschland, Verbraucherpreisindex (2005=100)".


Spain: Our data for Spain covers 18 provinces over 1995-2004 period. All series are from Spanish National Statistics Institute. For consumption we used "Household Final Consumption Expenditure by Autonomous Communities". Consumption is converted into per capita terms. For prices we used "Consumer Price Index by Autonomous Communities, Spain, Annual Means" with several base years, 1992, 2001, and 2006, which we converted into a common 2001 base year.
US: Data on the Gross Domestic Product (GDP) by state is available from the Bureau of Economic Analysis (BEA). Consumption series are not available at the state level for the U.S. As in Del Negro (2002), private consumption is calculated as total retail sales in state $i$ times the ratio of total retail sales to total consumption in the US. The data on total personal consumption expenditures is obtained from the NIPA Table published by the Bureau of Economic Analysis. Total retail sales numbers are published in Sales & Marketing Management’s Annual “Survey of Buying Power” (and reproduced in the Statistical Abstract of the United States). Up until the 1999 Survey, the data correspond to the previous calendar year. Starting in 2000, the Survey data correspond to the year in which they were published. This leads to a break in the data (missing data for 1999). CPI data is not available at the state level for the U.S. We use data from Del Negro (2002) up to 1995. From 1996 onwards, we follow del Negro’s procedure as closely as possible to extend the series to 2006.

A.2 Data summary statistics

In this Appendix we report the summary statistics for relative consumption and output growth and for the real exchange rate in our sample of countries. Table 11 reports the mean, standard deviation, minimum and maximum for the key variables of interest. These statistics are reported for the within-country pairs in the USA (Panel 1), Canada (Panel 2), Germany (Panel 3), Spain (Panel 4), and Japan (Panel 5); as well as for all cross-country pairs among Canada, US, Germany, Spain and Japan (Panel 5). Panel 5 also includes summary statistics on the components of the real exchange rate growth rate as shown in equation (2.5). These numbers show that international RER are significantly more volatile than RER for within-country pairs and are more volatile than international consumption differentials. Within-country RER are much smoother and in fact are

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52 Gross state product deflators currently in use are constructed taking into account differences in the product mix of states, but using national price data.
less volatile than relative consumption within our group of countries. Furthermore, as evident from Panel 5, the majority of cross-border RER fluctuations is due to movements in the nominal exchange rate.

Table 11: Sample summary statistics

<table>
<thead>
<tr>
<th>Panel</th>
<th>Variable</th>
<th>Obs</th>
<th>Mean (i)</th>
<th>Std. Dev. (ii)</th>
<th>Min (iv)</th>
<th>Max (v)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Panel 1: USA</td>
<td>Δ ln C_{i,t} - Δ ln C_{j,t}</td>
<td>41650</td>
<td>0.0001</td>
<td>0.0551</td>
<td>-0.4472</td>
<td>0.3859</td>
</tr>
<tr>
<td></td>
<td>Δ ln Y_{i,t} - Δ ln Y_{j,t}</td>
<td>41650</td>
<td>0.0000</td>
<td>0.0457</td>
<td>-0.3624</td>
<td>0.4277</td>
</tr>
<tr>
<td></td>
<td>Δ ln RER_{j;i}^{t}</td>
<td>41650</td>
<td>-0.0001</td>
<td>0.0117</td>
<td>-0.0922</td>
<td>0.0999</td>
</tr>
<tr>
<td>Panel 2: Canada</td>
<td>Δ ln C_{i,t} - Δ ln C_{j,t}</td>
<td>1674</td>
<td>0.0015</td>
<td>0.0182</td>
<td>-0.0680</td>
<td>0.1114</td>
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<tr>
<td></td>
<td>Δ ln Y_{i,t} - Δ ln Y_{j,t}</td>
<td>1674</td>
<td>0.0020</td>
<td>0.0557</td>
<td>-0.2988</td>
<td>0.4390</td>
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<tr>
<td></td>
<td>Δ ln RER_{j;i}^{t}</td>
<td>1674</td>
<td>0.0006</td>
<td>0.0095</td>
<td>-0.0331</td>
<td>0.0342</td>
</tr>
<tr>
<td>Panel 3: Germany</td>
<td>Δ ln C_{i,t} - Δ ln C_{j,t}</td>
<td>2040</td>
<td>0.0039</td>
<td>0.0183</td>
<td>-0.0515</td>
<td>0.0793</td>
</tr>
<tr>
<td></td>
<td>Δ ln Y_{i,t} - Δ ln Y_{j,t}</td>
<td>2040</td>
<td>0.0079</td>
<td>0.0410</td>
<td>-0.1696</td>
<td>0.1955</td>
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<td></td>
<td>Δ ln RER_{j;i}^{t}</td>
<td>2040</td>
<td>-0.0019</td>
<td>0.0182</td>
<td>-0.0890</td>
<td>0.0886</td>
</tr>
<tr>
<td>Panel 4: Spain</td>
<td>Δ ln C_{i,t} - Δ ln C_{j,t}</td>
<td>1377</td>
<td>-0.0035</td>
<td>0.0145</td>
<td>-0.0661</td>
<td>0.0473</td>
</tr>
<tr>
<td></td>
<td>Δ ln Y_{i,t} - Δ ln Y_{j,t}</td>
<td>1377</td>
<td>-0.0007</td>
<td>0.1364</td>
<td>-0.6737</td>
<td>0.6718</td>
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<td></td>
<td>Δ ln RER_{j;i}^{t}</td>
<td>1377</td>
<td>-0.0014</td>
<td>0.0046</td>
<td>-0.0171</td>
<td>0.0131</td>
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<tr>
<td>Panel 5: Japan</td>
<td>Δ ln C_{i,t} - Δ ln C_{j,t}</td>
<td>16215</td>
<td>0.0002</td>
<td>0.0251</td>
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<td>Δ ln Y_{i,t} - Δ ln Y_{j,t}</td>
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<td>0.0003</td>
<td>0.0224</td>
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<td>0.1033</td>
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<td>Δ ln RER_{j;i}^{t}</td>
<td>16215</td>
<td>-0.0002</td>
<td>0.0090</td>
<td>-0.0645</td>
<td>0.0713</td>
</tr>
<tr>
<td>Panel 6: All cross-country</td>
<td>Δ ln C_{i,t} - Δ ln C_{j,t}</td>
<td>93660</td>
<td>-0.0031</td>
<td>0.0367</td>
<td>-0.3250</td>
<td>0.3662</td>
</tr>
<tr>
<td></td>
<td>Δ ln Y_{i,t} - Δ ln Y_{j,t}</td>
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<td>0.0602</td>
<td>-0.7456</td>
<td>0.7908</td>
</tr>
<tr>
<td></td>
<td>Δ ln RER_{j;i}^{t}</td>
<td>93660</td>
<td>0.0070</td>
<td>0.0933</td>
<td>-0.2218</td>
<td>0.2272</td>
</tr>
<tr>
<td></td>
<td>Δ ln(P_{j;i} / P_{i;i})</td>
<td>93660</td>
<td>0.0085</td>
<td>0.0264</td>
<td>-0.1178</td>
<td>0.1270</td>
</tr>
<tr>
<td></td>
<td>Δ ln S_{i;j}^{t}</td>
<td>93660</td>
<td>-0.0015</td>
<td>0.0896</td>
<td>-0.2007</td>
<td>0.2007</td>
</tr>
</tbody>
</table>

Notes: The table reports summary statistics of the presented variables for all within country pairs in each country in our sample and all cross-country pairs (Panel 6). Obs. refer to the number of observations in each sample; Mean - sample average; Std. Dev. - sample standard deviation; Min-sample minimum; Max-sample maximum.
References


