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Keywords

Real exchange rate, risk sharing, border effect, intranational economics

JEL Classification

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Borders and Nominal Exchange Rates in Risk-Sharing*

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June 6, 2013

Abstract

Models of risk-sharing predict that relative consumption growth rates across locations should be positively related to real exchange rate growth rates across the same areas. 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 and show that it accounts for 53 percent 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? We find that over one-third of the border effect is due to nominal exchange rate fluctuations. We develop a simple open economy model that is consistent with the importance of nominal exchange rate 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

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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.¹

This paper revisits 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 rate relationship *within* countries. Our paper, to our knowledge, provides the first comprehensive evidence on this within-country relationship. By shedding light on the differences in the consumption-real exchange rate relationship within and between countries we may help to uncover the nature of the failure of full risk-sharing.

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. Since our methodology 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 in the Engel and Rogers (1996) study on deviations from the law of one price. In our framework, 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.² Following this logic, and using observed nominal exchange rates, we can decompose the border effect into that which can be 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, but not all of the border effect in the consumption-real exchange rate relationship. That is, even independent of fluctuations in bilateral exchange rate, consumption-real

¹See, for instance, Obstfeld (2007), Chari, Kehoe, and McGrattan (2002), Corsetti, Dedola, and Leduc (2008), Benigno and Thoenissen (2008), Kollmann (2012) and others.

²An 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.

exchange rate correlations across countries are lower than those across regions within countries.

Since the negative cross-border correlation of consumption and real exchange rates is to a large extent accounted for by the nominal exchange rate, could eliminating nominal exchange rate variability improve risk-sharing outcomes? This conclusion does not necessarily follow. Cross-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.³ 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. 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.

How much does the border contribute to the failure of consumption risk-sharing, taking account of both within and between country real exchange rate movements? We address this question by computing 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 percent of the deviation from full 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.

Can these findings be reconciled with standard models of real exchange rate determination? 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, particularly if financial markets are complete. On the other hand, many papers which offer potential resolutions to the Backus-Smith puzzle have no explicit role at all for the nominal exchange rate, so fixing the exchange rate would have no implications for the empirical tests of risk-sharing.⁴ More generally, although the presence

³This 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.

⁴In general, the literature discussing the Backus-Smith puzzle does not make any predictions about the behavior of the nominal exchange rate. See for instance Corsetti, Dedola, and Leduc (2008) and Benigno and Thoenissen (2008). 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

of incomplete markets and sticky prices may *facilitate* an explanation of the importance of the nominal exchange rate in the consumption-real exchange rate correlation, it is still not obvious what combination of shocks and propagation mechanisms could actually reproduce this phenomenon.

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 then ask whether this model exhibits the property that the sign of the consumption-real exchange rate correlation 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. The key insight that the model brings is that when the exchange rate is flexible, productivity shocks can lead the both the nominal and real exchange rate and relative consumption to move in different directions. By contrast, under a fixed exchange rate, the effect of the same shock on the real exchange rate is significantly dampened. In our baseline model, this difference is enough to reverse the sign of the 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.⁵

Thus, the role of the border in cross country consumption risk sharing is crucially tied to the nominal exchange rate regime, and the exchange rate displays characteristics in the model which are akin to those seen in the data.

As mentioned above, our analysis is related to an extensive recent 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 12 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

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.

⁵This 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.

real exchange rates.⁶ 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 $j = 1, \dots, J$ is given by:

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

where β is the subjective discount factor, $C_{j,t}$ denotes a composite consumption good in country j .⁷ 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 $S_t^{i,j}$ 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 $S_t^{i,j} = 1$. Then the real exchange rate between any two regions i and j in different countries is given by $RER_t^{j,i} = S_t^{i,j} P_{j,t} / P_{i,t}$, or $RER_t^{j,i} = P_{j,t} / P_{i,t}$ 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:

$$U_c(C_{i,t}) RER_t^{j,i} = U_c(C_{j,t}). \quad (2.1)$$

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 across countries. Then if utility is of a constant relative risk aversion (CRRA) form, with the

⁶Hess 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.

⁷It 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.

coefficient of relative risk aversion σ , 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:

$$\sigma(\Delta \ln C_{i,t} - \Delta \ln C_{j,t}) = \Delta \ln RER_t^{j,i}, \quad (2.2)$$

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).⁸

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, $\rho_{e,c_j/c_i} = \text{corr}(\Delta \ln RER_t^{j,i}, \sigma \Delta \ln \frac{C_{j,t}}{C_{i,t}})$, 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), Hess and Shin (1998), Del Negro (2002), Van Wincoop (1995), Crucini (1999), and others.

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.⁹ We employ intra-national data on consump-

⁸The estimation results with HP-filtered data are provided in the online appendix.

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)).

tion, 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 1995-2007 period, and 18 autonomous communities in Spain over 1995-2004 period.¹⁰ 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 following specification to link relative consumption growth and real exchange rate growth:

$$\Delta c_t^{i,j} = \beta_0 + \beta_1 \Delta e_t^{j,i} + \beta_2 (\Delta e_t^{j,i} * border_{i,j}) + v_t^{i,j},$$

where $v_t^{i,j}$ is the error term arising due to preference shocks, measurement error, etc. $border_{i,j}$ 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

$$\Delta c_t^{i,j} = \beta_0 + \beta_1 \Delta e_t^{j,i} + \beta_2 (\Delta e_t^{j,i} * border_{i,j}) + \beta_3 \Delta e_t^{j,i} * \ln \tilde{d}_{i,j} + v_t^{i,j}, \quad (2.3)$$

where $\ln \tilde{d}_{i,j}$ is the normalized log distance between any two locations i and j , defined as $\ln \tilde{d}_{i,j} = \ln d_{i,j} - \overline{\ln d_{i,j}}$. 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 $\overline{\ln d_{i,j}}$ is the average log distance between all cross-border pairs.¹¹ This normalization implies that $\ln \tilde{d}_{i,j}$ is equal to zero at $\ln d_{i,j} = \overline{\ln d_{i,j}}$, and simplifies interpretation of the β_2 coefficient, which now expresses the average effect of the

¹⁰These 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.

¹¹We measure distance in kilometers.

border for the consumption-real exchange rate relationship between any two locations that are $\ln \bar{d}_{i,j}$ kilometers away. The interaction term between the real exchange rate and distance allows the relationship between $\Delta c_t^{i,j}$ and $\Delta e_t^{j,i}$ to change monotonically with the distance.

Table 1: Estimates of Border Effect

	Pooled (i)	Fixed effects (ii)	Pooled (iii)	Fixed effects (iv)
$\Delta e_t^{j,i}$	0.349*** (0.019)	0.373*** (0.023)	0.275*** (0.019)	0.298*** (0.022)
$\Delta e_t^{j,i} * border_{i,j}$	-0.371*** (0.020)	-0.385*** (0.024)	-0.291*** (0.020)	-0.305*** (0.023)
$\Delta e_t^{j,i} * \ln \bar{d}_{i,j}$	0.021*** (0.003)	0.026*** (0.004)	0.019*** (0.003)	0.022*** (0.004)
$\Delta y_t^{i,j}$			0.169*** (0.006)	0.161*** (0.008)
$\Delta y_t^{i,j} * border_{i,j}$			-0.036*** (0.008)	-0.050*** (0.012)
$\Delta y_t^{i,j} * \ln \bar{d}_{i,j}$			0.040*** (0.004)	0.039*** (0.005)
$\Delta e_t^{j,i} + \Delta e_t^{j,i} * border_{i,j}$	-0.022*** (0.004)	-0.012*** (0.004)	-0.016*** (0.004)	-0.008** (0.004)
N	156509	156509	156509	156509

Notes: The dependent variable is relative consumption growth between locations i and j , $\Delta c_t^{i,j}$. The estimated specification in columns (i)-(ii) is equation (2.3); while in columns (iii)-(iv) it is equation (2.4). Robust standard errors are 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 1.¹² 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 countries 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. 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.¹³

How sensitive are these results to the assumption of complete access to capital markets? Many studies of risk-sharing, both intra-national and international, have relaxed this assumption 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

¹²In the fixed effects regression the fixed effects capture the time-invariant, bilateral pair specific jurisdiction effects.

¹³Note 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.

thumb, or in fact have no ability to borrow and lend at all.¹⁴ 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 $\Delta y_t^{i,j} = \Delta \ln Y_{i,t} - \Delta \ln Y_{j,t}$ 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 as follows:

$$\begin{aligned} \Delta c_t^{i,j} = & \beta_0 + \beta_1 \Delta e_t^{j,i} + \beta_2 (\Delta e_t^{j,i} * border_{i,j}) + \beta_3 \Delta e_t^{j,i} * \ln \tilde{d}_{i,j} \\ & + \beta_4 \Delta y_t^{i,j} + \beta_5 (\Delta y_t^{i,j} * border_{i,j}) + \beta_6 \Delta y_t^{i,j} * \ln \tilde{d}_{i,j} + v_t^{i,j} \end{aligned} \quad (2.4)$$

This specification allows for the border to affect the consumption-income relationship, and also for the the relationship between $\Delta c_t^{i,j}$ and $\Delta y_t^{i,j}$ 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 $\Delta c_t^{i,j}$ and $\Delta y_t^{i,j}$ in both pooled and fixed effects specifications. Allowing for limited participation also affects the within-country correlation between $\Delta c_t^{i,j}$ and $\Delta e_t^{j,i}$ 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 $\Delta c_t^{i,j}$ and $\Delta e_t^{j,i}$ becomes slightly, but significantly 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?

What explains the large drop in the consumption-real exchange rate correlation associated with crossing borders? An obvious fact 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

$$\Delta \ln RER_t^{j,i} = \Delta \ln(P_{j,t}/P_{i,t}) + \Delta \ln S_t^{i,j}, \quad (2.5)$$

where the first term on the right-hand-side captures movements in the 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.

¹⁴See for instance, Crucini (1999), Hess and Shin (2000), Hess and Shin (2010), Kollmann (2009) and Devereux, Smith, and Yetman (2009).

Table 2: Estimates of the Border Effect: RER Decomposition

	(i)	(ii)	(iii)	(iv)
$\Delta e_t^{j,i}$	0.384*** (0.023)	0.308*** (0.022)	0.387*** (0.023)	0.311*** (0.022)
$\Delta e_t^{j,i} * border_{i,j}$	-0.127*** (0.026)	-0.061** (0.025)	-0.137*** (0.026)	-0.068*** (0.025)
$\Delta e_t^{j,i} * \ln \tilde{d}_{i,j}$	0.038*** (0.004)	0.034*** (0.004)	0.042*** (0.004)	0.037*** (0.004)
$\Delta y_t^{i,j}$		0.159*** (0.008)		0.160*** (0.008)
$\Delta y_t^{i,j} * border_{i,j}$		-0.049*** (0.012)		-0.052*** (0.012)
$\Delta y_t^{i,j} * \ln \tilde{d}_{i,j}$		0.037*** (0.005)		0.038*** (0.005)
$\Delta \ln S_t^{i,j}$	-0.293*** (0.011)	-0.278*** (0.010)	-0.290*** (0.011)	-0.276*** (0.010)
$\Delta e_t^{j,i} * border_{i,j}^{SPA-GER}$			0.481*** (0.016)	0.375*** (0.017)
$\Delta e_t^{j,i} + \Delta e_t^{j,i} * border_{i,j}$	0.257*** (0.012)	0.248*** (0.010)	0.250*** (0.012)	0.242*** (0.010)
$\Delta e_t^{j,i} + \Delta e_t^{j,i} * border_{i,j} + \Delta \ln S_t^{i,j}$	-0.036*** (0.004)	-0.031*** (0.004)	-0.039*** (0.004)	-0.033*** (0.004)
$\Delta e_t^{j,i} + \Delta e_t^{j,i} * border_{i,j} + \Delta \ln S_t^{i,j} + \Delta e_t^{j,i} * border_{i,j}^{SPA-GER}$			0.441*** (0.015)	0.342*** (0.016)
N	156509	156509	156509	156509

Notes: The dependent variable is relative consumption growth between locations i and j , $\Delta c_t^{i,j}$. 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. Robust standard errors are 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 the negative border effect estimated in Table 1 can to a large extent 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 why the border effect would remain significant. Risk sharing within countries may be greater than that across countries due to deeper financial integration, a common language, common institutions, greater trade linkages and factor mobility, fiscal insurance mechanisms, etc. We investigate the quantitative role played by these factors in a related paper (Devereux and Hnatkovska (2011)). 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. Does this mean that eliminating nominal exchange rate flexibility would lead to an increase in the consumption-real exchange rate correlation? The problem with this interpretation 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. 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 3: Nominal Exchange Rate Variability

	US	Japan	Germany	Spain
Canada	1.3	2.9	2.6	2.6
US	-	2.7	2.3	2.3
Japan	-	-	2.9	3.1
Germany	-	-	-	0.6

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 percent. 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 $\Delta c_t^{j,i}$ 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.¹⁵ 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 $\Delta e_t^{j,i} * border_{i,j}^{GER-SPA}$ 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.¹⁶

To summarize, our findings suggest 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 percent and 80 percent in the case with and without hand-to-mouth consumers, respectively.

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 $\Delta c_t^{i,j}$ and $\Delta c_t^{j,i}$ to be country-specific. In particular, we amend the regression specification in equation (2.4) with the product terms of $\Delta e_t^{j,i}$ 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

¹⁵As 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.

¹⁶As 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.

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 $\Delta e_t^{j,i}$ 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 4: Estimates of the Border Effect: Accounting for country heterogeneity

	no $\Delta y_t^{i,j}$	with $\Delta y_t^{i,j}$
Panel (a). Homogeneous benchmark		
Intra-national	0.373*** (0.023)	0.298*** (0.022)
International	-0.012*** (0.004)	-0.008*** (0.004)
Panel (b). Heterogeneous benchmarks		
Intra-national correlations		
US	0.433*** (0.027)	0.329*** (0.027)
Canada	0.214*** (0.042)	0.159*** (0.045)
Germany	-0.201*** (0.021)	0.135*** (0.029)
Japan	0.360*** (0.034)	0.227*** (0.035)
Spain	0.925*** (0.144)	1.151*** (0.113)
International correlations		
US	-0.007** (0.004)	-0.005 (0.004)
Canada	-0.226*** (0.050)	-0.175*** (0.053)
Germany	-0.641*** (0.033)	-0.199*** (0.039)
Japan	-0.080* (0.043)	-0.107*** (0.043)
Spain	0.485*** (0.147)	0.817*** (0.116)

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 $\Delta e_t^{j,i}$ 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. Robust standard errors are in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively.

Table 4 reports our findings.¹⁷ 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 the benchmark for gauging the border effect (see columns (ii) and (iv) in Table 1). Panel (b) presents the estimation results that allow for country heterogeneity. We also report the results from the specification with hand-to-mouth consumers (columns labelled “with $\Delta y_t^{i,j}$ ”) and with homogeneous consumers (columns labelled “no $\Delta y_t^{i,j}$ ”).¹⁸ All regressions include jurisdiction-pairs fixed effects.

¹⁷Detailed estimation results are available from the authors upon request.

¹⁸The specification with the hand-to-mouth consumers also includes the interactive terms between output growth and dummy variables that identify within-country jurisdiction pairs for each country in our sample. This allows the consumption-output growth correlation to be country-specific as well.

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.¹⁹ 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.

Table 5: Estimates of Border Effect: Lagged consumption growth

	(i)	(ii)
$\Delta e_t^{j,i}$	0.370*** (0.032)	0.277*** (0.031)
$\Delta e_t^{j,i} * border_{i,j}$	-0.411*** (0.033)	-0.311*** (0.032)
$\Delta e_t^{j,i} * \ln \tilde{d}_{i,j}$	0.058*** (0.005)	0.052*** (0.005)
$\Delta y_t^{i,j}$		0.158*** (0.009)
$\Delta y_t^{i,j} * border_{i,j}$		-0.081*** (0.013)
$\Delta y_t^{i,j} * \ln \tilde{d}_{i,j}$		0.043*** (0.006)
$\Delta c_{t-1}^{i,j}$	-0.004 (0.003)	-0.013*** (0.003)
$\Delta e_t^{j,i} + \Delta e_t^{j,i} * border_{i,j}$	-0.040*** (0.005)	-0.034*** (0.005)
N	124588	124588

Notes: The dependent variable is relative consumption growth between locations i and j , $\Delta c_t^{i,j}$. 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. Robust standard errors are in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively.

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

¹⁹The only exception is Germany in the specification with no hand-to-mouth consumers, who exhibits negative intra-national consumption-RER correlation.

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. The results from the estimation are presented in Table 5.

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.²⁰

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, 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.²¹ We return to equation (2.2) and for each jurisdiction pair in our sample, for a choice of σ , we compute $\Delta c_t^{i,j} - (1/\sigma)\Delta e_t^{j,i}$ and obtain its standard deviation in the time-series.²² This gives us a cross-section of such standard deviations. If our choice of σ 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 6.

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

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

where $border_{i,j}$, is as before a dummy variable that takes on a value of one if locations i and

²⁰These results can be found in the online appendix to the paper.

²¹Engel 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.

²²We also conduct the analysis of this section using the mean of the squared difference, $mean \left[\Delta c_t^{i,j} - (1/\sigma)\Delta e_t^{j,i} \right]^2$, and find that the results are robust to this alternative measure of risk-sharing.

Table 6: Alternative measure of risk-sharing: Summary statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
(i) <i>std.dev.</i> $\left[\Delta c_t^{i,j} - (1/\sigma)\Delta e_t^{j,i} \right]$	10146	0.0514	0.0183	0.0039	0.1391
(ii) <i>std.dev.</i> $\left[\Delta c_t^{i,j} - (1/\sigma)\Delta \tilde{e}_t^{j,i} \right]$	10146	0.0332	0.0158	0.0039	0.1095

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.

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} .²³ Table 7 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. How important is the border? The average volatility of our risk-sharing measure is 5.14 percent, while the border coefficient is 2.74 percent, so the border accounts for 53 percent of the total, controlling for distance. Without the distance control, the contribution of the border is even larger at 77 percent.

Table 7: Alternative measure of risk-sharing: The role of the border

Variable	<i>border</i>	$\ln(\text{distance})$	N	explained by border
(i) <i>std.dev.</i> $\left[\Delta c_t^{i,j} - (1/\sigma)\Delta e_t^{j,i} \right]$	0.0395*** (0.0008)		10146	0.77
(ii) <i>std.dev.</i> $\left[\Delta c_t^{i,j} - (1/\sigma)\Delta e_t^{j,i} \right]$	0.0274*** (0.0009)	0.0090*** (0.0003)	10146	0.53
(iii) <i>std.dev.</i> $\left[\Delta c_t^{i,j} - (1/\sigma)\Delta \tilde{e}_t^{j,i} \right]$	0.0149*** (0.0008)		10146	0.45
(iv) <i>std.dev.</i> $\left[\Delta c_t^{i,j} - (1/\sigma)\Delta \tilde{e}_t^{j,i} \right]$	0.0113*** (0.0008)	0.0027*** (0.0002)	10146	0.34

Notes: This Table presents results from estimating regression specification (2.6). The dependent variable is noted in the first column of the table. Robust standard errors are in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively.

Can this border effect be attributed to movements in the nominal exchange rate, as argued previously? 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. But in the spirit of the previous regressions, we can decompose the real exchange rate into that component driven by relative price levels and

²³Note that this is a cross-sectional regression.

that by the nominal exchange rate. In particular, we make the following decomposition:

$$\Delta \ln RER_t^{i,j} = \Delta \ln(P_{j,t}\bar{S}_{i,j}/P_{i,t}) + \Delta \ln(S_t^{i,j}/\bar{S}_{i,j}), \quad (2.7)$$

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 variability²⁴. 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 $std.dev. [\Delta c_t^{i,j} - (1/\sigma)\Delta \tilde{e}_t^{j,i}]$, where $\Delta \tilde{e}_t^{j,i} = \ln(\widetilde{RER}_t^{j,i}) - \ln(\widetilde{RER}_{t-1}^{j,i})$.²⁵ Row (ii) of Table 6 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 7. 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. There we saw that controlling for the nominal exchange rate reduces the negative effect of the border on the consumption-real exchange rate correlation, but does not eliminate it.

How much does the border matter after we control for the nominal exchange rate in the calculation of relative prices? From panel (iv), the coefficient on the border dummy is 1.1 percent, while the average volatility of the adjusted risk-sharing measure as reported in Table 6 is 3.3 percent. So the border accounts for 34 percent 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 percent to 34 percent, 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 percent. These

²⁴Engel 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.

²⁵We 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.

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

From the previous section, it is apparent that the presence of the nominal 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 significantly 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.

What can account for these features? One potential reason that the nominal exchange rate affects risk sharing is that nominal goods prices are sticky, and so the stochastic properties of relative consumption and the real exchange rate differ between alternative exchange rate arrangements. But 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 property that the cross country risk sharing condition between consumption and real exchange rates holds exactly. On the other hand, many models in the literature which offer 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 would have no implications for the empirical tests of risk-sharing). More generally, although the presence of incomplete markets and sticky prices may *facilitate* an explanation of the importance of the nominal exchange rate in the consumption-real exchange rate correlation, it is still not obvious what combination of shocks and propagation mechanisms could actually reproduce this phenomenon.

In this section, we attempt to provide an account of the consumption-real exchange rate correlation by combining features of the previous literature on the Backus-Smith puzzle with a two country ‘New Keynesian’ model with gradual price adjustment. We then ask whether this model exhibits the property that the sign of the consumption-real exchange rate correlation depends upon the exchange rate regime.

Our model is essentially an extension of the recent literature on international real business cycles, in particular the model of Stockman and Tesar (1995), to an environment of sticky prices and alternative exchange rate regimes. The key insight that the model brings is that, when the exchange rate is flexible, productivity shocks can lead the both the nominal and real exchange rate and relative consumption to move in different directions. By contrast, under a fixed exchange rate, the effect of the same shocks on the real exchange rate is limited. In our baseline model, we show that this difference can be enough to reverse the sign of the consumption-real exchange rate correlation between exchange rate regimes. This result is especially highlighted when price

adjustment is more backward looking, as we show below.

3.1 A Model of the Consumption-Real Exchange Rate Correlation

Here we set out a two-sector model extended to a New Keynesian framework. 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. In each country consumers obtain utility from non-traded goods and a traded good aggregate. Traded goods aggregate consists of home and foreign traded goods. Home traded goods are also exported, and foreign goods are imported. We allow for shocks to productivity in the export and the non-traded goods sector. Let the utility of a representative infinitely lived home household evaluated from date 0 be defined as:

$$U_0 = E_0 \sum_{t=0}^{\infty} \beta^t \left(\frac{C_t^{1-\sigma}}{1-\sigma} - \frac{N_t^{1+\phi}}{1+\phi} \right), \quad \beta < 1, \quad (3.8)$$

where C_t is the composite home consumption bundle, and N_t is home labour supply. Composite consumption is defined as:

$$C_t = \left(\varrho^{\frac{1}{\varphi}} C_{Tt}^{1-\frac{1}{\varphi}} + (1-\varrho)^{\frac{1}{\varphi}} C_{Nt}^{1-\frac{1}{\varphi}} \right)^{\frac{\varphi}{1-\varphi}},$$

where C_{Tt} and C_{Nt} represent respectively, the composite consumption of traded and non-traded goods. The elasticity of substitution between traded and non-traded goods is φ , and the weight of tradable consumption in the aggregate consumption bundle is ϱ . Traded consumption in turn is decomposed into consumption of home goods, and foreign goods, as follows:

$$C_{Tt} = \left(\left(\frac{v}{2}\right)^{\frac{1}{\lambda}} C_{Ht}^{1-\frac{1}{\lambda}} + \left(1 - \frac{v}{2}\right)^{\frac{1}{\lambda}} C_{Ft}^{1-\frac{1}{\lambda}} \right)^{\frac{\lambda}{\lambda-1}},$$

where λ is the elasticity of substitution between home and foreign traded good and $v \geq 1$ captures a degree of home bias in preferences. In each case, we assume that consumption sub-aggregates are differentiated across the consumption of individual goods, with elasticity of substitution $\theta > 1$ across goods.

These consumption aggregates imply the following price index definitions:

$$P_t = \left(\varrho P_{Tt}^{1-\varphi} + (1-\varrho) P_{Nt}^{1-\varphi} \right)^{\frac{1}{1-\varphi}},$$

$$P_{Tt} = \left(\frac{v}{2} P_{Ht}^{1-\lambda} + \left(1 - \frac{v}{2}\right) P_{Ft}^{1-\lambda} \right)^{\frac{1}{1-\lambda}},$$

where P_{Tt} and P_{Nt} represent traded and non-traded price levels, and P_{Ht} and P_{Ft} are retail prices of home and foreign traded goods, respectively. P_t is the price of home aggregate consumption bundle. We define the real exchange rate as the relative price of the foreign consumption bundle

to the home bundle:

$$RER_t = \frac{S_t P_t^*}{P_t}.$$

Here S_t is the nominal exchange rate, and P_t^* is the price of foreign consumption bundle.

We assume that international financial markets are incomplete, using only non-contingent one period nominal bonds. The home budget constraint is given by:

$$P_t C_t + B_{t+1} + \frac{S_t B_{t+1}^*}{\Theta\left(\frac{S_t B_{t+1}^*}{P_t^*}\right)} = W_t N_t + \Pi_t + R_t B_t + R_t^* S_t B_t^*. \quad (3.9)$$

where B_{t+1} (B_{t+1}^*) indicates home (foreign) currency bond holdings of the home household, R_t (R_t^*) is the nominal interest rate on home (foreign) currency-denominated bonds, W_t represents the nominal wage, assumed to be equalized across sectors, and Π_t represents profits from the home firm plus net transfers from the government. The function $\Theta\left(\frac{S_t B_{t+1}^*}{P_t^*}\right)$ satisfies $\Theta\left(\frac{S_t B_{t+1}^*}{P_t^*}\right) = 1$, $\Theta' > 0$, $\Theta'' > 0$, and represents a transactions cost of foreign bond holdings, which is a 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 bond holdings in the usual way. The Euler equation for nominal bond pricing is given by:

$$\frac{C_t^{-\sigma}}{P_t} = R_{t+1} E_t \beta \frac{C_{t+1}^{-\sigma}}{P_{t+1}}. \quad (3.10)$$

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

$$W_t = C_t^\sigma N_t^\phi P_t \quad (3.11)$$

Foreign household's preferences and choices can be defined exactly symmetrically. The foreign representative household has weight $v/2$, $(1 - v/2)$ on the foreign (home) good in preferences over traded goods consumption.

Given a unified world market in bonds, the risk-sharing condition can be written as:

$$E_t \frac{C_{t+1}^{-\sigma}}{C_t^{-\sigma}} \frac{S_{t+1}}{S_t} \frac{P_t}{P_{t+1}} \Theta(\cdot) = E_t \frac{C_{t+1}^{*- \sigma}}{C_t^{*- \sigma}} \frac{P_t^*}{P_{t+1}^*} \quad (3.12)$$

This implies that up to a first-order approximation, expected consumption growth differentials across countries are positively related to expected changes in the real exchange rate (conditional on the small portfolio adjustment costs).

Firms in each sector produce using capital and labor. A typical firm in the non-traded (traded) sector has production function $Y_{Nt}(i) = A_{Nt} N_{Nt}(i)^{(1-\alpha)} K_{Nt}(i)^\alpha$, ($Y_{Ht}(i) = A_{Ht} N_{Ht}(i)^{(1-\alpha)} K_{Ht}(i)^\alpha$). Thus, there are two technology shocks - shocks to the non-traded sector A_{Nt} , and to the traded sector A_{Ht} . These shocks play substantially different roles in determining the consumption-real exchange rate correlation, as we argue below.

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

We assume that prices in each sector are sticky, and firms re-set prices according to a Calvo pricing policy, where the probability of re-adjusting price is sector specific, and is denoted as $1 - \kappa_j$ for sector j in each period. 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).²⁶ As we shall show below, this will 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 j of the home country is:

$$P_{j,t}^f(i) = \frac{E_t \sum_{z=0}^{\infty} m_{t+z} \kappa_j^z \frac{MC_{j,t+z}}{A_{j,t+z}} Y_{j,t+z}(i)}{E_t \sum_{z=0}^{\infty} m_{t+z} \kappa_j^z Y_{j,t+z}(i)}. \quad (3.13)$$

where m_t is a stochastic discount factor defined as $m_{t+z} = \beta^z \frac{C_{t+z}^{-\sigma}}{C_t^{-\sigma}}$, $Y_{j,t}(i)$ is demand for firm i 's good in sector j , and $MC_{j,t}$ is the nominal marginal cost.²⁷

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

$$P_{j,t}^b = \tilde{P}_{j,t-1} \frac{P_{t-1}}{P_t} \quad (3.14)$$

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

$$\tilde{P}_{j,t} = [(1 - \omega)P_{j,t}^{f(1-\theta)} + \omega P_{j,t}^{b(1-\theta)}]^{\frac{1}{1-\theta}} \quad (3.15)$$

In the aggregate, the price index for the home good in each sector then follows the process given

²⁶ Alternative approaches to introducing sluggishness in inflation response 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. We find that our results remain robust to these approaches.

²⁷ 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.

by:

$$P_{j,t} = [(1 - \kappa)\tilde{P}_{j,t}^{1-\theta} + \kappa P_{j,t-1}^{1-\theta}]^{\frac{1}{1-\theta}}. \quad (3.16)$$

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

Assume that the home country monetary authority follows a Taylor rule, which targets the CPI inflation rate $\Pi_t = P_t/P_{t-1}$. In addition, the monetary rule puts some weight on the nominal exchange rate. Specifically, the nominal interest rate R_{t+1} is set so that:

$$R_{t+1} = \beta^{-1} (\Pi_t)^{\gamma_\pi} (\Pi_{St})^\delta \quad (3.17)$$

where $\Pi_{St} = S_t/S_{t-1}$ as the gross rate of change of the nominal exchange rate. γ_π is the weight that monetary authority places on domestic inflation, and δ is the weight on exchange rate growth rate. For very high values of $\delta > 0$, the exchange rate will be effectively fixed.²⁸ The main issue of interest here will be to contrast the behavior of consumption and real exchange rates for values of $\delta = 0$, which we associate with a regime of inflation targeting in both countries, with a situation where the value of δ goes to a very high positive number, which we associated with a fixed exchange rate regime.

The foreign monetary authority follows a Taylor rule, but does not directly target the exchange rate.²⁹

Finally, goods market clearing conditions are given as:

$$\begin{aligned} Y_{Ht} &= C_{Ht} + C_{Ht}^* + I_{Ht} + I_{Nt}, \\ Y_{Ft}^* &= C_{Ft} + C_{Ft}^* + I_{Ft}^* + I_{Nt}^*, \\ Y_{Nt} &= C_{Nt}, \\ Y_{Nt}^* &= C_{Nt}^*. \end{aligned} \quad (3.18)$$

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:

$$E_t(\Delta c_{t+1} - \Delta c_{t+1}^*) = \frac{1}{\sigma} E_t \Delta r_{t+1} + \frac{1}{\sigma} \bar{\Theta} b_{t+1}^* \quad (3.19)$$

where $\Delta c_{t+1} = c_{t+1} - c_t$, $\Delta r_{t+1} = r_{t+1} - r_t$, and $\bar{\Theta}$ is a constant coefficient. 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

²⁸Note 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).

²⁹In fact, for the value of $corr(\Delta c - \Delta c^*, \Delta RER)$, 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 more easy to exposit.

small term, and has no first-order implication for the relationship between relative consumption and real exchange rates.

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 consumption growth and real exchange rate changes. Define $y_t \equiv \Delta c_{t+1} - \Delta c_{t+1}^*$, and $x_t \equiv \frac{1}{\sigma} \Delta rer_{t+1}$. Then, ignoring the bond market adjustment term, the risk sharing condition can be stated as

$$y_{t+1} = x_{t+1} + \nu_{t+1} \quad (3.20)$$

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 (3.20) is

$$\hat{\beta}_{x,y} = 1 + \frac{\text{Cov}(x_{t+1}, \nu_{t+1})}{\text{Var}(x_{t+1})} \quad (3.21)$$

This can be negative only if innovations to relative consumption growth less real exchange rate growth, ν_t , covary negatively with the real exchange rate. As pointed out by Benigno and Thoenissen (2008), productivity shocks to the traded goods sector, in an environment of incomplete financial markets, can generate such a negative coefficient. The key is in the behavior of various components of the real exchange rate.

The real exchange rate in the model can be decomposed into a function of three relative prices:

$$rer_t = \rho(\tau_{Nt}^* - \tau_{Nt}) + (1 - \rho)(v - 1)\tau_t + \rho\tau_t \quad (3.22)$$

Here $\tau_{Nt} = p_{Nt} - p_{Ht}$, $\tau_{Nt}^* = p_{Nt}^* - p_{Ft}^*$, and τ_t is the terms of trade, defined as $s_t + p_{Ft}^* - p_{Ht}$. The real exchange rate then moves for two reasons. First, it is affected by cross country differences in internal relative prices. If the foreign country relative price of non-traded goods rises, relative to the home country relative price, there will be a home real depreciation. Secondly, the real exchange rate is affected by the external terms of trade τ_t . For a given value of $\tau_{Nt}^* - \tau_{Nt}$, a terms of trade deterioration causes a real exchange rate depreciation, since it raises the foreign relative to the domestic price of consumption, and does this more so, the greater the degree of home bias in preferences.

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, matching the properties of the US economy in annual 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 8.

The discount factor is set at 0.99. The labor supply elasticity ϕ is set at unity. The elasticity of intertemporal substitution is set at 0.5 so that $\sigma = 2$. Under a fixed exchange rate, δ is set at a high value so that the nominal exchange rate is constant (the actual value for δ is irrelevant once it is set high enough). We set $\gamma_\pi = 1.5$. The elasticity of substitution between non-traded and traded

Table 8: Benchmark model parameters

	parameter	value
subjective discount factor	β	0.99
labor supply elasticity	ϕ	1
risk aversion	σ	2
elasticity of substitution b/n N and T goods	φ	0.5
elasticity of substitution b/n H and F goods	λ	2
T goods weight in aggregate consumption	ϱ	0.6
home bias in T goods sector	v	1.25
depreciation rate	δ_K	0.025
capital share	α	0.33
1-probability of re-adjusting price in sector N	κ_N	0.8
1-probability of re-adjusting price in sector T	κ_T	0.67
weight on inflation in interest rate rule	γ_π	1.5
weight on exchange rate growth in interest rate rule	δ	0, or $\lim \delta \rightarrow \infty$
fraction of rule of thumb firms	ω	0.3
investment adjustment costs	$\frac{\phi'' i/k}{\phi}$	-0.75
bond adjustment costs	Θ	0.0025

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 $\lambda = 2$, a benchmark estimate.³⁰ In the analysis below we also show how the results are affected by variations in λ . We assume that the non-traded goods sector is 60 percent of GDP, somewhat smaller than that of the US economy, but larger than that for smaller economies. This pins down parameter ϱ . 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 percent 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 $\frac{\phi'' i/k}{\phi} = -0.75$. 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 $\kappa_N = 0.8$ and $\kappa_T = 0.67$. 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 ω to match the estimates of Galí and Gertler (1999), who find that the backward looking coefficient is around 0.3. This implies that about a quarter of firms in each sector use a rule of thumb approach to price setting rather than a forward looking approach.

In measuring the persistence and volatility of productivity shocks to the traded and non-traded sector, we follow Benigno and Thoenissen (2008). They estimate home traded sector productivity shocks to have persistence $\mu_H = 0.85$ and standard deviation $\sigma_H = 1.94$ percent, while non-traded

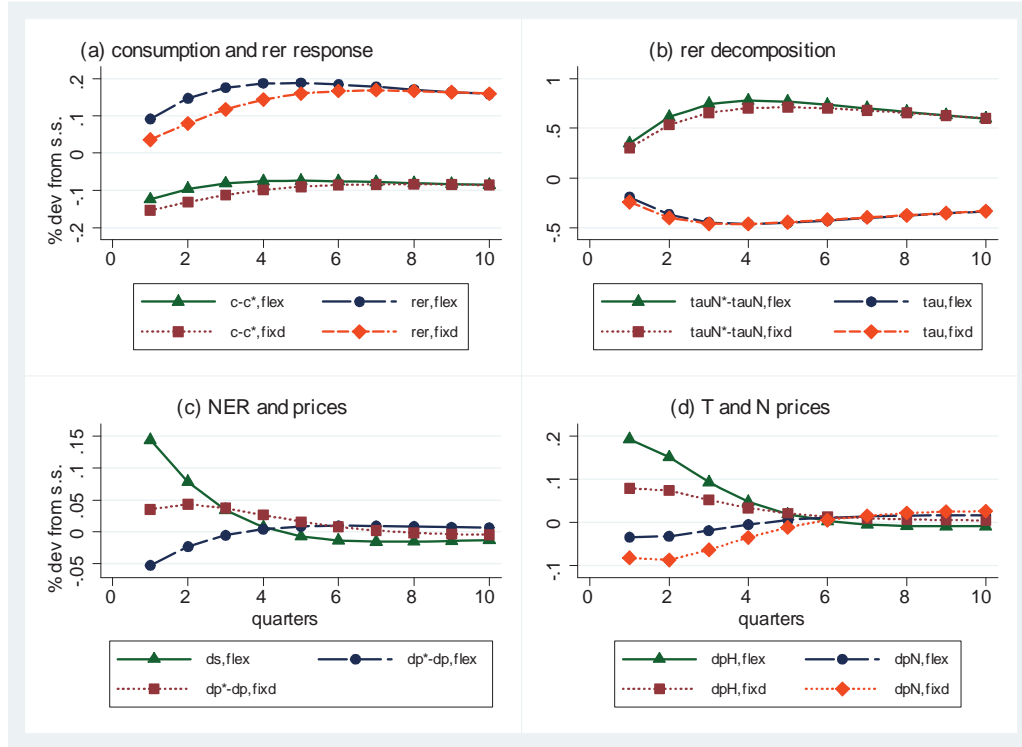
³⁰See Ruhl (2005), Matsumoto (2007) among others.

shocks are much less persistent and less volatile, with $\mu_N = 0.3$ and $\sigma_N = 0.71$. The foreign productivity process is symmetric. In addition, we impose the covariance structure of shocks as estimated by Benigno and Thoenissen (2008).³¹

3.3 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. Figure 2 reports the corresponding impulse responses following a negative shock to non-traded sector productivity.

Figure 1: Impulse responses following a negative shock to T sector productivity



Notes: The figures present the impulse responses of home country's relative consumption, 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.

³¹In particular, taking the shock innovations at $\varepsilon = [\varepsilon_H, \varepsilon_F^*, \varepsilon_N, \varepsilon_N^*]'$, we use $Var(\varepsilon) = \begin{bmatrix} 3.76 & 1.59 & 0.72 & 0.44 \\ 1.59 & 3.76 & 0.44 & 0.72 \\ 0.72 & 0.55 & 0.51 & 0.21 \\ 0.44 & 0.72 & 0.21 & 0.51 \end{bmatrix}$.

Panel (a) of Figures 1 and 2 illustrates the responses of home relative consumption and real exchange rate depreciation. 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 τ , the home terms of trade (as shown in equation (3.22)). Panel (c) shows the alternative decomposition of the real exchange rate, breaking down the response between the nominal exchange rate growth, Δs , 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 Δp_H and the home non-traded good Δp_N . Both Figures contrast the dynamics under a Taylor rule ('flex') with those under an exchange rate peg ('fixd').

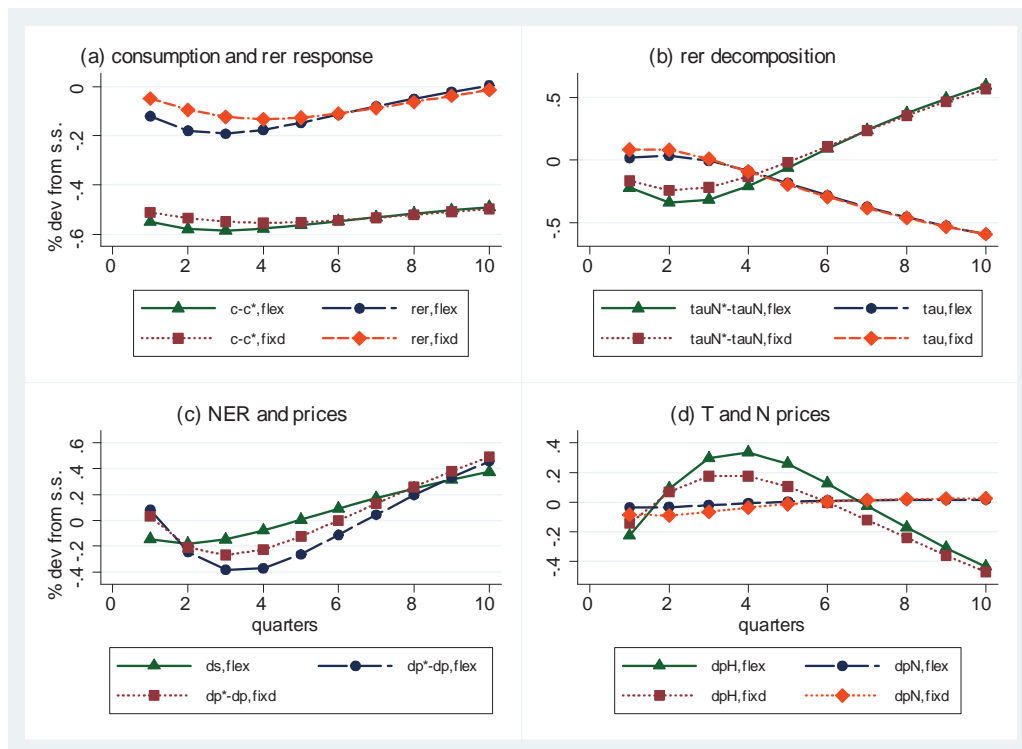
Panel (a) of Figure 1 establishes that in both cases – Taylor rule policy and fixed exchange rate – the decline in home traded goods productivity leads to a fall in relative consumption, and a real exchange rate depreciation. The real depreciation is achieved by a rise in $\tau_N^* - \tau_N$, the 'relative-relative' 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 τ_N , implying a real exchange rate depreciation. At the same time, there is a fall in τ , 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 τ , so the net effect amounts to a real exchange rate depreciation.

How do 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, the nominal exchange rate immediately depreciates. 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, the real depreciation must be fully 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 more deflation, 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 stifled, while under a Taylor rule, the immediate response of the nominal exchange rate facilitates a much larger immediate response in the real exchange rate. As we see below, in the simulated model this difference may be big enough to account for the differences in the consumption-real exchange rate correlations across alternative exchange rate regimes.

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 τ . 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.

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



Notes: The figures present the impulse responses of home country's relative consumption, 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, due to the immediate response of the nominal exchange rate to a traded goods productivity shock, this negative correlation may be higher under a Taylor rule policy than under a fixed exchange rate. In order to explore this question, we turn to model simulations.

Table 9 reports the simulation results for the model under benchmark calibration. The Table summarizes the values of $\text{corr}(\Delta c - \Delta c^*, \Delta rer)$, the standard deviation of $\Delta c - \Delta c^*$, the real exchange rate, output, investment, and the nominal exchange rate, under the case of a Taylor rule, and under the exchange rate peg. 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 $\text{corr}(\Delta c - \Delta c^*, \Delta rer)$ is -0.03 , approximately that we found in the data for floating exchange rate pairs of regions. By contrast, under a fixed exchange rate, the correlation is 0.11 . Thus, we find that the value of $\text{corr}(\Delta c - \Delta c^*, \Delta rer)$, in the model does depend on the nominal monetary rule. The correlation is positive under fixed exchange rates, and negative under flexible exchange rates. This is consistent with the empirical findings above.

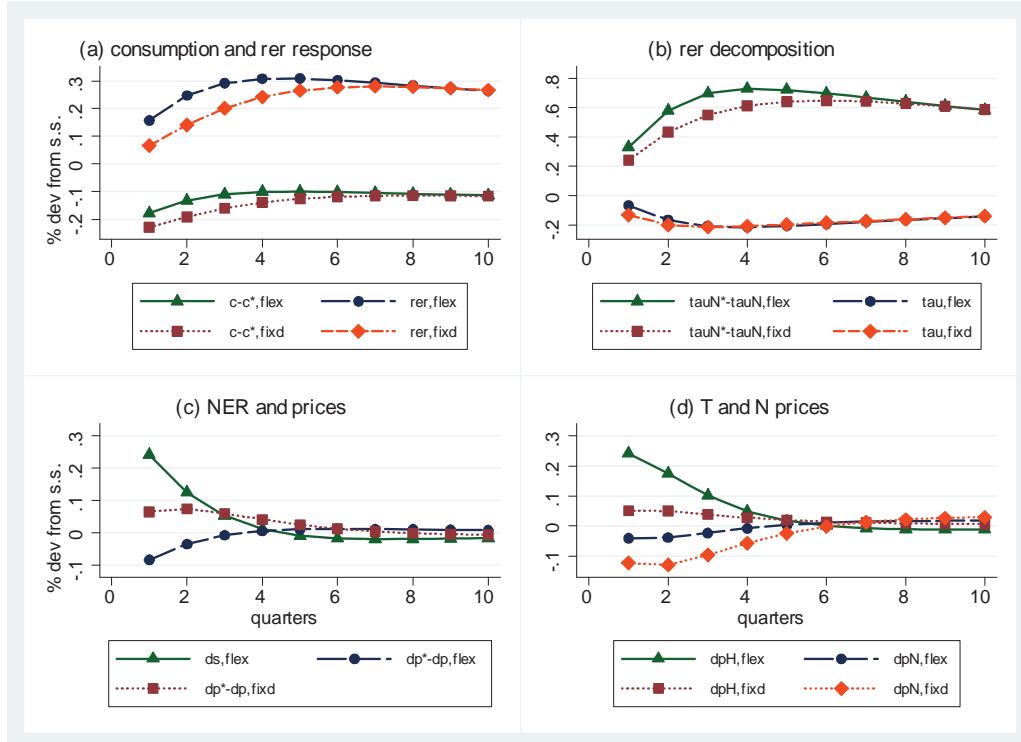
Table 9: Simulation results

Policy	$\text{stdev}(\Delta c - \Delta c^*)$	$\text{stdev}(\Delta rer)$	$\text{corr}(\Delta c - \Delta c^*, \Delta rer)$	$\text{stdev}(y)$	$\text{stdev}(I)$
$\delta = 0$	0.88	0.76	-0.03	1.64	4.46
$\lim \delta \rightarrow \infty$	0.96	0.59	0.11	1.73	4.5

How robust is this finding? 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 3 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.

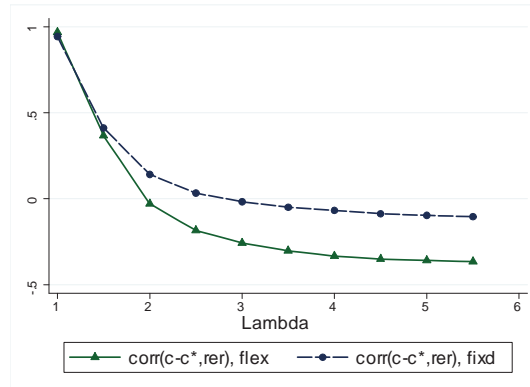
Figure 4 illustrates the results for the consumption-real exchange rate correlation under a fixed and flexible exchange rate regimes, as a function of the size of the trade elasticity λ . With λ close to unity, the correlation is equal to unity, since we have effectively complete markets in that case (see Cole and Obstfeld (1991)). As λ increases, we find that the correlation declines and is systematically lower for a flexible exchange rate policy than under fixed exchange rates. For values of λ in the range of 2 to 3, we see that the correlation reverses sign with the exchange rate regime. As λ 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 3: Impulse responses following a negative shock to T sector productivity with $\lambda = 5$



Notes: The figures present the impulse responses of 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.

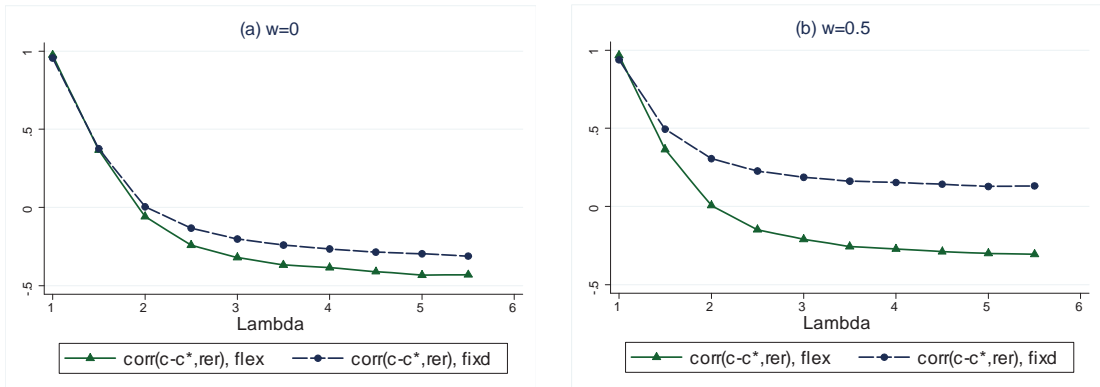
Figure 4: Consumption-RER correlation: Sensitivity to trade elasticity



Notes: The figure presents a plot of C-RER correlation in the model for various values of parameter λ . Correlations are presented under both inflation targeting (flex) and fixed (fixd) exchange rate regimes.

How important is the assumption of backward looking prices, implied by the parameter ω ? Figure 5, panel (a) repeats Figure 4 but now assuming that $\omega = 0$, so that there are no rule of thumb price setters. Here again, we see the negative relationship between $\text{corr}(\Delta c - \Delta c^*, \Delta rer)$ and λ still prevails. Moreover, the correlation is still more negative under a Taylor rule than in a fixed exchange rate. But in contrast to the baseline case, the correlation is significantly negative for all empirically reasonable values of λ , for both the Taylor rule and the pegged exchange rate. Intuitively, in the absence of rule of thumb 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 λ than in the baseline model.

Figure 5: Consumption-RER correlation: Sensitivity to backward-looking price setting



Notes: The figure presents a plot of C-RER correlation in the model for various values of parameter λ . Correlations are presented under both inflation targeting (flex) and fixed exchange rate regimes (fixd). Panel (a) presents the results with parameter $\omega = 0$, while panel (b) is with $\omega = 0.5$.

Finally, panel (b) of Figure 5 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 Figure 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 $\lambda \geq 2$ always negative under a flexible exchange rate.

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 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.

References

- ABEL, A. B. (1990): "Asset Prices under Habit Formation and Catching Up with the Joneses," *American Economic Review*, 80(2), 38–42.
- ARELLANO, M., AND S. BOND (1991): "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations," *Review of Economic Studies*, 58(2), 277–97.
- ASDRUBALI, P., B. SØRENSEN, AND O. YOSHA (1996): "Channels of Interstate Risksharing: United States 1963-1990," *Quarterly Journal of Economics*, 111, 1081–1110.
- ATHANASOULIS, S. G., AND E. VAN WINCOOP (2001): "Risk Sharing Within the United States: What do Financial Markets and Fiscal Federalism Accomplish?," *The Review of Economics and Statistics*, 83(4), 688–698.
- BACKUS, D. K., AND G. W. SMITH (1993): "Consumption and Real Exchange Rates in Dynamic Economies with Non-traded Goods," *Journal of International Economics*, 35(3-4), 297–316.
- BAYOUMI, T., AND M. W. KLEIN (1997): "A Provincial View of Economic Integration," *IMF Staff Papers*, 44(4), 534–556.
- BENIGNO, G., P. BENIGNO, AND F. GHIRONI (2007): "Interest rate rules for fixed exchange rate regimes," *Journal of Economic Dynamics and Control*, 31(7), 2196–2211.

- BENIGNO, G., AND C. THOENISSEN (2008): “Consumption and Real Exchange Rates with Incomplete Markets and Non-traded Goods,” *Journal of International Money and Finance*, 27(6), 926–948.
- CHARI, V. V., P. J. KEHOE, AND E. R. MCGRATTAN (2002): “Can Sticky Price Models Generate Volatile and Persistent Real Exchange Rates?,” *Review of Economic Studies*, 69(3), 533–63.
- CHRISTIANO, L. J., M. EICHENBAUM, AND C. L. EVANS (2005): “Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy,” *Journal of Political Economy*, 113(1), 1–45.
- CLARIDA, R., J. GALI, AND M. GERTLER (2002): “A simple framework for international monetary policy analysis,” *Journal of Monetary Economics*, 49(5), 879–904.
- COLE, H. L., AND M. OBSTFELD (1991): “Commodity trade and international risk sharing : How much do financial markets matter?,” *Journal of Monetary Economics*, 28(1), 3–24.
- CORSETTI, G., L. DEDOLA, AND S. LEDUC (2008): “International Risk Sharing and the Transmission of Productivity Shocks,” *Review of Economic Studies*, 75(2), 443–473.
- CRUCINI, M. (1999): “On International and National Dimensions of Risksharing,” *The Review of Economics and Statistics*, 81(1), 73–84.
- DEL NEGRO, M. (2002): “Asymmetric Shocks Among U.S. States,” *Journal of International Economics*, 56(2), 273–297.
- DEVEREUX, M. B., AND C. ENGEL (2003): “Monetary Policy in the Open Economy Revisited: Price Setting and Exchange-Rate Flexibility,” *Review of Economic Studies*, 70(4), 765–783.
- DEVEREUX, M. B., AND V. HNATKOVSKA (2011): “International and Intra-national Real Exchange Rates: Theory and Evidence,” Working paper, University of British Columbia.
- DEVEREUX, M. B., G. W. SMITH, AND J. YETMAN (2009): “Consumption and Real Exchange Rates in Professional Forecasts,” NBER Working Papers 14795, National Bureau of Economic Research, Inc.
- ENGEL, C. (1999): “Accounting for U.S. Real Exchange Rate Changes,” *Journal of Political Economy*, 107(3), 507–538.
- ENGEL, C., AND J. ROGERS (1996): “How wide is the border?,” *The American Economic Review*, 86(5), 1112–1125.
- GALI, J., AND M. GERTLER (1999): “Inflation Dynamics: A Structural Econometric Analysis,” *Journal of Monetary Economics*, 44, 195–222.
- GALI, J., M. GERTLER, AND J. D. LOPEZ-SALIDO (2001): “European inflation dynamics,” *European Economic Review*, 45(7), 1237–1270.

- GHIRONI, F., AND M. J. MELITZ (2005): “International Trade and Macroeconomic Dynamics with Heterogeneous Firms,” *The Quarterly Journal of Economics*, 120(3), 865–915.
- GORODNICHENKO, Y., AND L. TESAR (2009): “Border effect or country effect? Seattle may not be so far from Vancouver after all,” *American Economic Journal: Macroeconomics*, 1(1), 219–241.
- HADZI-VASKOV, M. (2008): “Does the Nominal Exchange Rate Explain the Backus-Smith Puzzle? Evidence from the Eurozone,” Working Papers 07-32, Utrecht School of Economics.
- HESS, G., AND K. SHIN (1998): “Intranational Business Cycles in the United States,” *Journal of International Economics*, 44, 289–314.
- HESS, G., AND E. VAN WINCOOP (2000): *Intranational Macroeconomics*. Cambridge: Cambridge University Press.
- HESS, G. D., AND K. SHIN (2000): “Risk Sharing by Households Within and Across Regions and Industries,” *Journal of Monetary Economics*, 45(3), 533–560.
- HESS, G. D., AND K. SHIN (2010): “Understanding the Backus-Smith puzzle: It’s the (nominal) exchange rate, stupid,” *Journal of International Money and Finance*, 29(1), 169–180.
- KOLLMANN, R. (1995): “Consumption, Real Exchange Rates and the Structure of International Asset Markets,” *Journal of International Money and Finance*, 14(2), 191–211.
- KOLLMANN, R. (2009): “Household Heterogeneity and the Real Exchange Rate: Still a Puzzle,” CEPR Discussion Papers 7301, C.E.P.R. Discussion Papers.
- (2012): “Limited asset market participation and the consumption-real exchange rate anomaly,” *Canadian Journal of Economics*, 45(2), 566–584.
- MANKIW, N., AND R. REIS (2002): “Sticky information versus sticky prices: a proposal to replace the New Keynesian Phillips curve,” *The Quarterly Journal of Economics*, 117(4), 1295–1328.
- MATSUMOTO, A. (2007): “The Role of Nonseparable Utility and Nontradeables in International Business Cycle and Portfolio Choice,” IMF Working Papers 07/163, International Monetary Fund.
- MCADAM, P., AND A. WILLMAN (2004): “Supply, Factor Shares and Inflation Persistence: Re-examining Euro-area New-Keynesian Phillips Curves,” *Oxford Bulletin of Economics and Statistics*, 66(s1), 637–670.
- MENDOZA, E. G. (1995): “The Terms of Trade, the Real Exchange Rate, and Economic Fluctuations,” *International Economic Review*, 36(1), 101–37.
- MUSSA, M. (1986): “Nominal exchange rate regimes and the behavior of real exchange rates: Evidence and implications,” *Carnegie-Rochester Conference Series on Public Policy*, 25(1), 117–214.

- NAKAMURA, E., AND J. STEINSSON (2008): “Five Facts about Prices: A Reevaluation of Menu Cost Models,” *The Quarterly Journal of Economics*, 123(4), 1415–1464.
- OBSTFELD, M. (2007): “International Risk Sharing and the Costs of Trade,” The ohlin lectures, Stockholm School of Economics.
- OBSTFELD, M., AND K. ROGOFF (1996): *Foundations of International Macroeconomics*. MIT Press Cambridge, MA.
- OKAWA, Y., AND E. VAN WINCOOP (2010): “Gravity in International Finance,” Working Papers 072010, Hong Kong Institute for Monetary Research.
- PORTES, R., AND H. REY (2005): “The determinants of cross-border equity flows,” *Journal of International Economics*, 65(2), 269–296.
- RAVN, M. O. (2001): “Consumption Dynamics and Real Exchange Rate,” CEPR Discussion Papers 2940, C.E.P.R. Discussion Papers.
- RUHL, K. (2005): “The elasticity puzzle in international economics,” *unpublished, University of Texas at Austin*.
- SCHMITT-GROHE, S., AND M. URIBE (2003): “Closing small open economy models,” *Journal of International Economics*, 61(1), 163–185.
- STOCKMAN, A. C. (1987): “The Equilibrium Approach to Exchange Rates,” *Economic Review, Federal Reserve Bank of Richmond*, 12(30).
- STOCKMAN, A. C., AND L. L. TESAR (1995): “Tastes and Technology in a Two-Country Model of the Business Cycle: Explaining International Comovements,” *American Economic Review*, 85(1), 168–185.
- VAN WINCOOP, E. (1995): “Regional Risksharing,” *European Economic Review*, 39, 1545–1567.

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)".

Japan: Our data for Japan covers 47 prefectures over 1990-2005 period. All series are from "The Annual Report on Prefectural Accounts" published by ESRI. Consumption series is "Private household and nonprofit consumption, real, constant prices 2000". For prices we used "GDP Deflator: Expenditure, Constant Price 2000" series. Consumption data is converted into per capita terms.

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.³² 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.

³²Gross state product deflators currently in use are constructed taking into account differences in the product mix of states, but using national price data.

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 10 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 6). Panel 6 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 less volatile than relative consumption within our group of countries. Furthermore, as evident from Panel 6, the majority of cross-border RER fluctuations is due to movements in the nominal exchange rate.

Table 10: Sample summary statistics

Variable	Obs (i)	Mean (ii)	Std. Dev. (iii)	Min (iv)	Max (v)
Panel 1: USA					
$\Delta \ln C_{i,t} - \Delta \ln C_{j,t}$	41650	0.0001	0.0551	-0.4472	0.3859
$\Delta \ln Y_{i,t} - \Delta \ln Y_{j,t}$	41650	0.0000	0.0457	-0.3624	0.4277
$\Delta \ln RER_t^{j,i}$	41650	-0.0001	0.0117	-0.0922	0.0999
Panel 2: Canada					
$\Delta \ln C_{i,t} - \Delta \ln C_{j,t}$	1674	0.0015	0.0182	-0.0680	0.1114
$\Delta \ln Y_{i,t} - \Delta \ln Y_{j,t}$	1674	0.0020	0.0557	-0.2988	0.4390
$\Delta \ln RER_t^{j,i}$	1674	0.0006	0.0095	-0.0331	0.0342
Panel 3: Germany					
$\Delta \ln C_{i,t} - \Delta \ln C_{j,t}$	2040	0.0039	0.0183	-0.0515	0.0793
$\Delta \ln Y_{i,t} - \Delta \ln Y_{j,t}$	2040	0.0079	0.0410	-0.1696	0.1955
$\Delta \ln RER_t^{j,i}$	2040	-0.0019	0.0182	-0.0890	0.0886
Panel 4: Spain					
$\Delta \ln C_{i,t} - \Delta \ln C_{j,t}$	1377	-0.0035	0.0145	-0.0661	0.0473
$\Delta \ln Y_{i,t} - \Delta \ln Y_{j,t}$	1377	-0.0007	0.1364	-0.6737	0.6718
$\Delta \ln RER_t^{j,i}$	1377	-0.0014	0.0046	-0.0171	0.0131
Panel 5: Japan					
$\Delta \ln C_{i,t} - \Delta \ln C_{j,t}$	16215	0.0002	0.0251	-0.1310	0.1614
$\Delta \ln Y_{i,t} - \Delta \ln Y_{j,t}$	16215	0.0003	0.0224	-0.0946	0.1033
$\Delta \ln RER_t^{j,i}$	16215	-0.0002	0.0090	-0.0645	0.0713
Panel 6: All cross-country					
$\Delta \ln C_{i,t} - \Delta \ln C_{j,t}$	93660	-0.0031	0.0367	-0.3250	0.3662
$\Delta \ln Y_{i,t} - \Delta \ln Y_{j,t}$	93660	-0.0004	0.0602	-0.7456	0.7908
$\Delta \ln RER_t^{j,i}$	93660	0.0070	0.0933	-0.2218	0.2272
$\Delta \ln(P_{j,t}/P_{i,t})$	93660	0.0085	0.0264	-0.1178	0.1270
$\Delta \ln S_t^{i,j}$	93660	-0.0015	0.0896	-0.2007	0.2007

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.