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Real Exchange Rates, Preferences, and Incomplete Markets: Evidence, 1961-2001

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Abstract
Many international macroeconomic models link the real exchange rate to a ratio of marginal utilities. We examine this link empirically, allowing the marginal utility of consumption to depend on government expenditure, real money balances, or external habit. We also consider two environments with incomplete asset markets; one with exogenously missing markets but an endogenous discount rate that anchors the distribution of wealth and one with endogenous market segmentation. Although none of these satisfies theoretical and over-identifying restrictions for every country, utility with external habit persistence provides the best match with real exchange rates for OECD countries between 1961 and 2001.

Keywords: real exchange rate, consumption, marginal utility

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1. Introduction

A large volume of research in international macroeconomics focuses on explanations for the persistence in real exchange rates. Most models of persistent deviations from purchasing power parity imply that relative marginal utilities of consumption across countries should be as persistent as the real exchange rate. Unfortunately, there is little evidence of this link between the real exchange rate and relative consumption. For example, work by Chari, Kehoe and McGratten (2000) replicates the autocorrelation of the real exchange rate, but the autocorrelation of relative consumption in their model is almost twice the empirical level. Obstfeld and Rogoff (2000) list this 'disconnect' among the key, unresolved issues in open-economy macroeconomics.

This paper re-examines this issue in historical data, using a variety of marginal utility models. These augmented models include government spending, leisure, real money balances, and an external habit stock, under a variety of functional forms. We seek to identify a utility model that resolves the inconsistency of the predicted real exchange rate with its empirical counterpart. We also consider two models of incomplete asset markets. In one, non-contingent bonds are the only asset, while an endogenous discount rate anchors the distribution of wealth. In the other, markets are endogenously segmented to a degree that depends on the rate of inflation.

Roughly speaking, traditional models of marginal utility imply that a country undergoing a real depreciation also should experience relatively rapid consumption growth, with the scale between the two changes governed by the elasticity of intertemporal substitution in consumption. The logic can be understood by considering a positive, country-specific, supply shock, which tends to raise national consumption and lower national prices at the same time. But the empirical evidence in this study suggests that the consumption change may be in the opposite direction, and that real depreciations are weakly associated with relatively slow consumption growth.

However, this effect is statistically insignificant. Real exchange rates tend to be quite persistent, but their growth rates are not closely linked to those of current, relative consumption. We find evidence that rates of real depreciation *are* related to a moving average of relative consumption growth. This finding can be interpreted as evidence of external habit persistence, as introduced by Abel (1990). The implication is that these preferences might be worth adopting in open-economy, macroeconomic models.

The remainder of this paper is organized as follows. Section 2 provides background on the condition tested in this paper. Section 3 outlines a parametric model of utility, allowing a role for government spending. Section 4 gives statistical evidence for OECD countries since 1961. Section 5 extends the estimating equations to reflect in turn, leisure, real balances, and external habit persistence. Section 6 examines the models with incomplete asset markets. Section 7 concludes.

2. Estimating Equations

Obstfeld (1989), Backus and Smith (1993), and Kollmann (1995) observed that a range of international macroeconomic models with non-traded goods link the real exchange rate to relative consumption. Suppose that two countries i and j have price levels p_i and p_j. Let the nominal exchange rate between them, e_{ij}, be the local currency price for one unit of country j's currency in terms of country i's currency. Suppose that utility U depends on a vector of quantities x that includes consumption c, with marginal utility U_c. Then under complete asset markets the real exchange rate is equal to the ratio of marginal utilities of consumption:

$$\frac{e_{ijt}\ p_{jt}}{p_{it}} = \frac{U_c(x_{jt})}{U_c(x_{it})} \tag{1}$$

As noted by Obstfeld (1989), Apte, Sercu and Uppal (2001), and Engel (2000), this link between relative prices and relative quantities holds under complete asset markets even if there are frictions in goods markets, including non-traded goods, pricing to market (PTM), local currency pricing (LCP), or transport costs. Obstfeld (1994, theorem) and Apte, Sercu, and Uppal (2001, proposition 2) provide comprehensive derivations of this condition. This paper reconsiders several parametric models of utility used in studies which embody this condition. We use the method-of-moments links between the endogenous variables to estimate preference parameters without specifying a complete model and to test over-identifying restrictions . General equilibrium models including these preferences will have a chance to fit the dynamics of consumption and exchange rates only if they pass this test.

If this static condition holds in levels, then it also holds in growth rates. Define Δ as the *gross* growth rate operator, so that $\Delta x_t = x_t/x_{t-1}$. Then:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \frac{\Delta U_c(x_{jt})}{\Delta U_c(x_{it})}. \tag{2}$$

This version, in gross growth rates, holds that *ex post*, intertemporal, marginal rates of substitution are equal. It is convenient for statistical inference, because the growth rates are often stationary. This stochastic singularity (2) obviously will not hold in historical data, so we examine the conditional forecast:

$$\mathrm{E}_t\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \mathrm{E}_t\frac{\Delta U_c(x_{jt})}{\Delta U_c(x_{it})}. \tag{3}$$

While condition (3) is clearly weaker than (2) it is a necessary condition, and so useful for testing. As we shall see, it typically provides enough information to reject (2) statistically. It also is consistent with preference shocks, provided these are not persistent.

Empirical results are presented for balanced panels – with preference parameters common across countries – and for country-pairs, with the US in each case acting as the reference country j. Choosing instruments z_{it} for country i amounts to asking which shocks are expected to be insured. Instruments include lagged endogenous variables – such as relative consumption growth or lagged residuals – which presumably reflect many shocks. But we also could instrument using exogenous variables like policy changes or natural disasters. The moment conditions we adopt for estimation are:

$$\mathrm{E}\left[z_{it-1}\left(\Delta\frac{e_{ijt}p_{jt}}{p_{it}} - \frac{\Delta U_c(x_{jt})}{\Delta U_c(x_{it})}\right)\right] = 0. \tag{4}$$

Estimation employs iterated GMM. For some country pairs the growth in relative consumption does not display strong persistence in the data, which limits the choice of relevant instruments as discussed by Stock and Wright (2000). For this reason, the instrument set includes lagged residuals, which reflect the differential persistence in rates of real depreciation and relative consumption growth.

The findings in this paper update and extend those of Obstfeld (1989, 1994), Backus and Smith (1993), Kollmann (1995), and Apte, Sercu, and Uppal (2001) who studied isoelastic utility with marginal utility depending only on current consumption. Obstfeld, Kollmann, and Apte, Sercu, and Uppal also were concerned to formulate the problem so that least-squares regression methods could be used, whereas GMM is applied directly in this exercise. Ravn (2001), in a related study, also uses instrumental variable estimation, but on a logarithmic approximation to (3). He considers a different range of possible models of marginal utility. In addition, this paper augments utility with external habit persistence as applied by Abel (1990, 1999) and Campbell and Cochrane (1999) to explain several features of asset prices. Finally, models of incomplete asset markets have been considered by Obstfeld (1989) and Kollmann (1995) to explain the real exchange rate. They showed that trade only in non-contingent bonds equalizes the expected, intertemporal marginal rates of substitution, so that again equation (3) holds, though (2) does not. We extend their work by considering utility functionals with stochastic discount rates that are consistent with a stationary distribution of wealth when markets are exogenously incomplete. We also examine the implications of recent research on endogenously segmented markets.

3. Benchmark Utility Model

Suppose that the discount factors are constant and equal. Period utility in country i is of the power form:

$$u(x_{it}) = \begin{cases} x_{it}^{1-\alpha}/(1-\alpha) & \alpha > 0, \quad \alpha \neq 1; \\ \ln x_{it} & \alpha = 1 \end{cases} \tag{5}$$

where x_{it} is an aggregator over private consumption c_{it} and government consumption g_{it}. This aggregate in turn is of the CES form:

$$x_{it} = [\mu c_{it}^{\omega} + (1-\mu) g_{it}^{\omega}]^{1/\omega}. \tag{6}$$

We examine two special cases of this aggregator. In the first, $\omega = 0$, which yields the Cobb-Douglas case:

$$x_{it} = c_{it}^{\mu} g_{it}^{1-\mu}. \tag{7}$$

In the second special case, $\mu = 1$ so that public expenditure does not directly affect utility and $x_{it} = c_{it}$.

With this parametric utility model (5) and (6), the estimating equations (2) become:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{1-\omega}\left(\frac{\Delta x_{it}}{\Delta x_{jt}}\right)^{\alpha+\omega-1}, \tag{8}$$

where the CES functional form is used for utility. The Cobb-Douglas special case, with $\omega = 0$, gives:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{1-\mu(1-\alpha)}\left(\frac{\Delta g_{it}}{\Delta g_{jt}}\right)^{-(1-\mu)(1-\alpha)}. \tag{9}$$

Finally, the traditional case where $\mu = 1$ and $\omega = 0$, so that utility depends only on private consumption is given by:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{\alpha}. \tag{10}$$

We also consider a second aggregator:

$$x_{it} = c_{it} + \psi g_{it}, \tag{11}$$

as used by Christiano and Eichenbaum (1992) in a business-cycle model. This functional form leads to the estimating equations:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it} + \psi g_{it}}{\Delta c_{jt} + \psi g_{jt}}\right)^{\alpha}. \tag{12}$$

4. Statistical Results

Data are for a set of ten countries: Canada, Denmark, Finland, France, Italy, Japan, New Zealand, Sweden, the UK, and the US. This group was selected based on the availability of measures for private consumption excluding durables. The data run from the 1960s for Canada, the UK and the US, from the 1970s for Denmark, Finland, France, and Italy, and from the 1980s for Japan, Sweden and New Zealand. The data reflect recent changes in accounting practices in many countries. For instance, the US National Accounts are now constructed according to the chain-weighted accounting standard. The appendix provides exact data definitions and sources.

Estimates are based on a balanced panel from 1981:II to 1999:III. The panel excludes Denmark and New Zealand so as to include relatively long time spans for the remaining countries. We also provide estimates of country-specific preference parameters using the estimating equations (and all data) for individual countries. In all cases, the reference country j is the United States. This is a natural choice because the U.S. data span is greater than that for any other country. In any case, the panel results are not sensitive to the choice of reference country, for the GMM estimator – like GLS – takes into account the correlation between residuals that may be caused by a U.S. shock.

The data include some observations from the Bretton Woods period of fixed exchange rates for the U.S, UK, and Canada. The properties of real exchange rates tend to differ across nominal exchange-rate regimes. We do not exclude these observations from the results for individual countries, for if the theory is a useful guide then the corresponding ratio of marginal utilities also should have time series properties that vary across monetary regimes.

Results from the CES functional form for utility specified by equation (8) are not reported because the three preference parameters were not readily identifiable, and $\hat{\omega}$ was not significant. As a result, the model is further restricted with $\omega = 0$ to give the Cobb-Douglas specification with private consumption and government spending as in (9). Table 1 contains the results. The parameters include the curvature of utility and the share of private consumption in utility, $\hat{\mu}$. The first row provides results from the panel. The instrument set consists of a constant and once-lagged, own residuals. The estimated coefficient of relative risk aversion, $\hat{\alpha}$, is negative, while the consumption share is a small, positive fraction. Neither is statistically significant (at conventional levels), despite the large number of instruments.

Table 1 also shows the results of estimation country-by-country. The parameters are over-identified with four instruments, which include a constant, growth in lagged relative consumption, lagged relative government spending, and the lagged residual. The estimated coefficient of relative risk aversion, $\hat{\alpha}$, is negative for seven of the nine countries. In the remaining two countries – Japan and Finland – the estimate is positive but smaller than its standard error. Estimates of the share of consumption in utility, $\hat{\mu}$, are within a range centered around unity, where six of nine estimates are significantly different from zero. The p-values for the J-test of over-identifying restrictions range from 1% to 34%, and in five of the nine cases yield rejections at the 5% level of significance.

When we considered the alternative aggregator (11) over consumption and government spending we again could not find a significant role for government spending in the marginal utility model. Estimates of α were generally negative, while those of ψ were statistically insignificant.

Table 2 specializes the utility function with the restriction $\mu = 1$ which yields equation (10), the gross-growth rate version of the relationship studied by Backus and Smith (1993) and Kollmann (1995). Table 2 suggests that there is no statistical relationship between the growth in this measure of marginal utility and the real exchange rate. The panel estimate of α is again an insignificant, negative number. The p-value for the J-test is 0.39, so that the over-identifying restrictions cannot be rejected (at conventional levels). The picture that emerges is that the difference between the growth in the real exchange rate and the growth in relative consumption scaled by α is essentially unpredictable – as required by theory – but involves a value of α that is zero or slightly negative.

In the remaining rows of table 2, the estimated coefficients $\hat{\alpha}$ also are negative for five of the nine country pairs, contradicting the concavity of utility. Only for Italy is $\hat{\alpha}$ positive and significant at conventional levels of significance. The p-values for the J-test of the over-identifying restrictions range from 2% to 22%, and yield rejections at the 5% level in about half the cases.

We also examined the possibility of weak identification, as studied by Stock and Wright (2000). First, as already noted, by including a lagged residual in the instrument set in table 2 (and other tables) we make identification stronger than if lagged, relative consumption growth were the only instrument. Second, we also estimated α using only a constant term as an instrument, since in that case the parameter clearly is strongly identified. The results (not shown) were very similar to those in table 2: six point estimates were negative, and no estimate was positive and larger than its estimated standard error. Third, we examined the sensitivity of the findings to the choice of weighting matrix, since weak identification is a function of the combination of this matrix, the moment condition, and the instrument set. Again the results did not change. We concluded that the findings probably are not due to weak identification but rather to the choice of functional form or to the risk-sharing condition's failure.

To summarize, there is little evidence to support the real exchange rate model when utility is defined over private consumption and government spending. Although the J-test statistics sometimes are small, the estimated coefficients of relative risk aversion are often negative and insignificant.

5. Extending the Utility Models

Given the weak results so far, the next step is to examine other models of marginal utility that have been used in international macroeconomics. These models include utility functions that are nonseparable between private consumption and either real money balances or leisure. In addition, environments with external habit persistence have proved successful at replicating aspects of asset prices and business cycles. Most of these models imply that a multiplicative factor – in relative employment, money, or lagged consumption – is missing from the benchmark, risk-sharing condition (10). The omission of these variables might therefore account for the negative findings from tables 1 and 2.

5.1 Leisure

First, consider a nonseparability between consumption and leisure. The period utility function is again of power form:

$$u(d_{it}) = \begin{cases} d_{it}^{1-\alpha}/(1-\alpha) & \alpha > 0, \quad \alpha \neq 1; \\ \ln d_{it} & \alpha = 1. \end{cases} \tag{13}$$

where d_{it} combines the CES aggregate x_{it}, (6) of private and public consumption with a measure of employment l_{it}:

$$d_{it} \equiv x_{it} - \delta l_{it}^{\eta}. \tag{14}$$

This form is chosen to nest the case studied by Greenwood, Hercowitz and Huffman (1988), Hercowitz and Sampson (1991), Devereux, Gregory, and Smith (1991), and Correia, Neves, and Rebelo (1995). Those authors set $\alpha = 1$ and $x_{it} = c_{it}$ for analytical tractability. Because we are not solving a model, we consider a more general form.

As an example, consider the case in which $x_{it} = c_{it}$, as we found in section 4. Then the estimating equations are:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left[\frac{\Delta\left(c_{it} - \delta l_{it}^{\eta}\right)}{\Delta\left(c_{jt} - \delta l_{jt}^{\eta}\right)}\right]^{\alpha}. \tag{15}$$

The relative growth of marginal utility across countries now contains labour supply measures. International employment differences tend to be positively autocorrelated, so this addition to the statistical model may produce persistence that more closely matches the persistence in the real exchange rate.

Equations (15) are estimated with l_{it} measured as the OECD's index of total employment. The results were entirely negative. Parameters often could not be identified, and for those country-pairs yielding results, the estimates $\hat{\delta}$ had the wrong sign and $\hat{\alpha}$ was negative. As a result, the evidence does not provide support for the time series correlation between relative employment and relative prices to match the restrictions of this marginal utility model.

5.2 Money

A number of recent research papers in international finance have included real money balances in the utility function. We next follow Chari, Kehoe, and McGrattan (2000) who incorporate a nonseparability between real balances, m_{it}, and private consumption. Their period utility function is:

$$u(c_{it}, m_{it}) = \frac{1}{1-\alpha}\left[\left(\mu c_{it}^{\omega} + (1-\mu)m_{it}^{\omega}\right)^{\frac{1}{\omega}}\right]^{1-\alpha}, \tag{16}$$

which gives rise to the following marginal utility with respect to consumption:

$$u_c = \mu c_{it}^{\omega-1}\left[\mu c_{it}^{\omega} + (1-\mu)m_{it}^{\omega}\right]^{\frac{1-\alpha-\omega}{\omega}}. \tag{17}$$

Real balances are measured as broad money (generally M3) divided by the consumption deflator. The equations for the real exchange rate are estimated using both the CES and Cobb-Douglas aggregators over private consumption and real balances. The CES model does not yield statistically significant parameter estimates for $\hat{\omega}$. The necessary condition for the specialized, Cobb-Douglas version is:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{1-\mu(1-\alpha)}\left(\frac{\Delta m_{it}}{\Delta m_{jt}}\right)^{-(1-\mu)(1-\alpha)}. \tag{18}$$

Results are shown in table 3. In the panel, the estimated weight on consumption is $\hat{\mu} = 0.867$ so that the weight on real balances is 0.133. This value is estimated with some precision and seems plausible given the theory. The curvature of the utility function also now is clearly significant (at conventional levels), but again is negative.

For individual countries, the estimated curvature of utility, $\hat{\alpha}$, is negative for all but the Japan–US case, and standard errors are larger than the estimates for all but the New Zealand–US case. There is some evidence that μ is a fraction – implying that real balances join consumption in the utility function – but only for New Zealand is $\hat{\mu}$ estimated precisely and below 1 at conventional significance levels. In seven of the nine cases, the J-test rejects the over-identifying restrictions at the 5% level. Overall, including real money balances in the model of marginal utility does not lead to a statistical improvement or to interpretable preference parameters.

5.3 External Habit

The final extension of the utility model adopts external habit persistence. External habit has proved successful in explaining aspects of the equity premium puzzle. The utility model follows Abel (1990, 1999). The period utility function is related to a benchmark level of utility, s_{it}, in country i that is treated as exogenous to the numerous identical households:

$$u(c_{it}, s_{it}) = \frac{1}{1-\alpha}\left(\frac{c_{it}}{s_{it}}\right)^{1-\alpha}. \tag{19}$$

The benchmark utility level is

$$s_{it} = C_{it}^{\delta_0} C_{it-1}^{\delta_1} (\xi_i^t)^{\delta_2} \tag{20}$$

where C is aggregate consumption, $\xi_i \geq 1$ so that the benchmark consumption level grows exogenously over time, and the other parameters satisfy $0 \leq \delta_0, \delta_1, \delta_2 \leq 1$. Not all of the preference parameters in (19) and (20) are identifiable, as Abel (1999) discusses. Specifically, define ϕ, θ, and κ_{ij} as:

$$\begin{aligned} \phi &= \alpha - \delta_0(\alpha - 1) > 0, \\ \theta &= \delta_1(\alpha - 1) \\ \kappa_{ij} &= \left(\frac{\xi_i}{\xi_j}\right)^{\delta_2(1-\alpha)} \end{aligned} \tag{21}$$

which are linear combinations of the underlying preference parameters. Applying the external habit persistence model to an international environment then yields the following relationship between consumption and the real exchange rate,

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{\phi}\left(\frac{\Delta c_{it-1}}{\Delta c_{jt-1}}\right)^{-\theta}\kappa_{ij}, \tag{22}$$

where κ_{ij} reflects the potentially different growth rates of reference utility.

Table 4 contains the results from estimating equation (22). We focus first on the panel results. The point estimates $\hat{\phi}$ and $\hat{\theta}$ imply various, plausible combinations of preference parameters via the restrictions (21). For example, $\alpha = 3$, $\delta_0 = 0.58$, and $\delta_1 = 0.76$ are possible. In addition, $\hat{\kappa}$ is precisely estimated and indistinguishable from 1, so that the benchmarks for consumption grow at the same rate in all countries. Finally, the J-test statistic has a p-value of 0.85, indicating a non-rejection of the over-identifying restrictions.

These results stand in contrast to those for individual countries, where in each case both preference parameters are not significantly different from zero at conventional levels of significance. Notice that the panel estimates do not appear to be weighted averages of those from individual countries; that is because the panel uses a different sample so as to include all but two countries. Thus it omits some Japanese and UK data, for example.

According to the habit-persistence model, the benchmark model with consumption alone (equation (10) and table 2) may lead to an inconsistent estimator of α because the expression for marginal utility omits lagged consumption growth. Comparing the panel results in tables 2 and 4 shows that the estimated preference parameters are significant, and accord with theory, once lagged consumption is included in the model of marginal utility.

6. Incomplete Asset Markets

Several studies of international business cycles have worked with exogenously incomplete asset markets, in which there is trade only in non-contingent bonds. This approach also breaks the period-by-period connection between the real exchange rate and relative consumption; the real exchange rate deviates from the stochastic singularity given by equation (2) because financial assets do not span all contingencies. But, without further modification, such incomplete asset markets imply a non-stationary distribution of wealth across countries. As a result, a stable equilibrium is no longer well defined. An endogenous discount factor that evolves stochastically alleviates this problem. As explained and derived by Lucas and Stokey (1984), the endogenous discount rate increases marginal 'impatience' as the economy accumulates net foreign assets, so that the distribution of wealth evolves along a stationary path.

The utility model incorporating an endogenous discount rate is Schmitt-Grohé's and Uribe's (2003) model 1a, which uses modified Uzawa (1968) preferences. This model includes power utility in consumption and a discount factor that also depends on consumption:

$$\beta(c_{it}) = \big(1 + c_{it}\big)^{-\gamma}, \qquad \gamma \geq 0 \tag{23}$$

so that impatience rises as consumption rises. Importantly, households do not internalize the fact that the discount factor depends on consumption. Alternately, imagine that β depends on per capita consumption which the household takes as exogenous. Schmitt-Grohé and Uribe show that an open-economy model with this feature behaves identically to one in which households internalize the effect of consumption on impatience. It also behaves very similarly to other models of incomplete markets with stationary wealth distributions, such as those with a debt-elastic, international, interest-rate differential.

In aggregate data with power sub-utility, marginal utility then is:

$$U_{ci} = \big(1 + c_{it}\big)^{-\gamma} c_{it}^{-\alpha}. \tag{24}$$

The real exchange rate condition with an endogenous discount rate is:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{\alpha} \left[\frac{\Delta(1 + c_{it-1})}{\Delta(1 + c_{jt-1})}\right]^{\gamma}. \tag{25}$$

The results for the endogenous discount model are reported in table 5. In the panel, the coefficient of relative risk aversion is 1.53, with a small standard error. The J-test cannot reject the overidentifying restrictions, as the p-value is 0.73. As in table 4, therefore, the addition of higher-order dynamics in consumption allows us to identify a positive value for $\hat{\alpha}$. In this case, though, the additional parameter γ is estimated to be a significant, negative number, which is inconsistent with the theory (23), wherein impatience rises with consumption.

In the results for individual countries relative to the U.S., standard errors again are too large to allow us to draw conclusions about the parameters. Also noteworthy is the observation that the p-values for the J-test tend to be quite low (4 of 9 are less than 0.10) which suggests that adding instruments to try to improve precision might well lead to rejection of the moment restrictions.

Our final extension of the estimating equation involves a change in the constraints, rather than the utility function. Alvarez, Atkeson, and Kehoe (2002) study a cash-in-advance economy in which households are subject to an additional cost to transfer money in or out of the asset market. With this friction, they show that the real exchange rate is equal to the ratio of international marginal utilities of households that are active in asset markets. Consumption by those households is denoted with the subscript a so that:

$$\frac{e_{ijt}p_{jt}}{p_{it}} = \frac{U_c(x_{ajt})}{U_c(x_{ait})}. \tag{26}$$

This condition differs from the one we have studied so far, for it depends not on aggregate consumption but rather on the consumption of households that are active in asset markets.

Condition (26) is not directly testable because the x_{ait} are not observed, and so Alvarez, Atkeson, and Kehoe discuss ways to test it indirectly. For example, they study the implications of this endogenous market incompleteness for the correlation between real and nominal exchange rates, using data for high-inflation countries such as Turkey, Israel, Mexico, and several countries in Latin America. One implication of their modelling of endogenous market incompleteness is that the consumption of active households will be more closely approximated by aggregate consumption when the inflation rate is high. In fact, at high enough inflation rates all households will be active and the usual condition (1) will hold. Unfortunately, measures of consumption excluding durables (or even total consumption for long time spans) typically are unavailable for high-inflation countries. Consequently, we explore the implications of endogenous segmentation for the OECD real exchange rates in our original sample.

Denote the inflation rate in country i by π_{it}. Then the revised condition is:

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{\alpha}\left(1 - \exp[-\lambda(\pi_{it} + \pi_{jt})]\right). \tag{27}$$

Thus if both inflation rates are zero, the added term is is 0 and there is no connection between the real exchange rate and the ratio of marginal utilities measured in aggregate consumption. Negative inflation rates are not observed in our sample. In contrast, as the inflation rates rise the new, weighting term goes to one at a rate estimated via the parameter λ. In that case, there will be a connection between the real exchange rate and the ratio of marginal utilities, as predicted by the simplest case of the benchmark theory. The revised estimating equations (27) imply that the standard versions should fit better during periods when both countries have high inflation – such as the 1970s – than during periods of shared, low inflation – such as the 1990s.

The results (not shown in a table) did not support equations (27) for these OECD countries. Estimates of λ were positive, but estimates of α remained generally negative. Thus the addition of this weighting factor did not yield a readily interpretable relation between real exchanges rates and measured, marginal utility. Perhaps this market segmentation holds in less-developed countries with higher inflation rates, but consumption data typically are not available for long time spans for those countries.

7. Conclusion

This paper extends empirical work investigating the link between relative, international marginal utilities and the real exchange rate. We study models of marginal utility that include government spending, leisure, real balances, or external habit. We also allow for incomplete asset markets with a stochastic discount rate or for endogenous market segmentation.

Country-by-country estimation generally does not yield precise estimates of the preference parameters, while panel estimation does so. Real exchange rates tend to be quite persistent, so that their growth rates are difficult to predict. Consumption tends to roughly follow a random walk so that its growth rate also is difficult to predict. Thus identifying the parameters is difficult without international panels.

In the panels, in turn, the results are negative with one, conspicuous exception. The model with external habit yields significant coefficients, with signs in accord with theory. For example, one identification gives a coefficient of relative risk aversion of 3, a weight of 0.58 on current consumption, and a weight of 0.76 on lagged consumption. This specification of marginal utility also passes the J-test of over-identification. It implies that the benchmark model including only current consumption may suffer from omitted-variables bias. Our brief investigation thus serves as a first step, before going on to study a fully-solved model with external habit.

Recently, Corsetti, Dedola, and Leduc (2001) and Duarte and Stockman (2001) have modelled real exchange rates in environments with *both* goods market segmentation and incomplete asset markets. While these two studies have many distinct features, they each find that the combination of these two market imperfections potentially can break the link between the real exchange rate and relative consumption. Testing these real exchange rate models requires measuring shocks to productivity and money growth, and so these models do not lend themselves to direct estimation by GMM. Nevertheless, the calibrated example of Corsetti, Dedola, and Leduc shows that the correlation between the real exchange rate and relative consumption can be zero or negative, as we found in tables 1 and 2 for example. Historical sample paths predicted from these models remain to be studied. Our negative results on many of the alternatives reinforce the idea that these models with two frictions also are worthy of further investigation.

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Data Appendix

The data for real government expenditure, g, real private consumption, c, and the implicit price deflator of private consumption, p are from the OECD *Quarterly National Accounts*. The data are generally constructed using the fixed-weight standard of the 1993 System of National Accounts. The US data, however, are chain-weighted series. Base years vary by country, but are generally set consistent with 1992 constant prices. The reference year for the US chain-weighted data is 1996. Nominal exchange rates, e, come from the IMF's *International Financial Statistics* (IFS) database. Exchange rates are local currency prices of a US dollar, and reflect the average rate over the quarter. Private consumption in constant prices is measured as the sum of consumer nondurables and services expenditure in constant prices. For the US case, the consumption variable is calculated as the sum of nondurables and services expenditure in current prices, deflated by the implicit price deflator for final consumption. This procedure for the US consumption variable is adopted because the disaggregated chain-weighted series are not additive across the components.

Labour data are non-seasonally adjusted index series with 1995 acting as the base year. These data are generally defined as 'total employment' in each country, so that individuals who have worked for any duration over the reference period are included as employed. The quarterly series are collected from the *Main Economic Indicators*, published by the OECD.

Broad money data are also extracted from the *International Financial Statistics* database. These series are generally defined as M3, but for some countries the only available data are M2 or M4. Specifically, M3 is available for Denmark, Finland, France, New Zealand, Sweden and the US. Data for Japan and the UK are reported as M4. Finally, broad money is reported as M2 for Canada and Italy. Real money balances are defined as the nominal money supply in local currency divided by the consumption price deflator.

The panel includes Canada, Denmark, Finland, France, Italy, Japan, Sweden, the United Kingdom, New Zealand and the United States. The panel composition was selected based on the availability of disaggregated consumption data for nondurables and services expenditure. Graphs of the series were examined to identify possible problems with the data. Seasonal patterns in Swedish government spending and consumption, in Japanese consumption, the employment index data, and the broad money data for Denmark, New Zealand and Sweden were removed with the `esmooth` command in Rats.^{TM} Table A1 gives the time span available for each time series.

Table A1
Availability of Time Series

	Canada	Denmark	Finland	France	Italy
c	61:1–00:1	88:1–00:4	75:1–00:4	70:1–98:4	70:1–98:3
g	61:1–00:4	88:1–00:4	75:1–00:4	78:1–00:4	70:1–00:4
p	61:1–00:4	88:1–00:4	75:1–00:4	78:1–00:4	70:1–00:4
e	55:1–01:4	55:1–01:4	55:1–01:4	55:1–01:4	55:1–01:4
l	60:1–00:4	80:1–01:4	60:1–01:1	64:4–00:4	60:1-00:4
M	68:1–00:4	70:1–00:4	75:1–98:4	75:1–98:4	75:1-98:4

	Japan	Sweden	UK	US	New Zealand
c	70:1–99:1	80:1–98:4	55:1–00:4	59:1–00:4	83:1–00:2
g	80:1–00:4	80:1–98:4	55:1–99:1	59:1–00:4	82:2–00:2
p	80:1–00:4	80:1–98:4	55:1–00:4	59:1–00:4	82:2–00:2
e	55:1–01:1	55:1–01:1	55:1–01:1	–	82:1–01:1
l	60:1-01:1	61:2-01:1	78:2-00:4	60:1-01:1	60:1–01:1
M	80:1-00:4	62:1-00:4	63:1-00:4	62:1-00:4	65:4–00:4

Table 1
Power/Cobb-Douglas Utility in Consumption and Government Spending

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{1-\mu(1-\alpha)} \left(\frac{\Delta g_{it}}{\Delta g_{jt}}\right)^{-(1-\mu)(1-\alpha)}$$

Country	T	$\hat{\alpha}$ (s.e.)	$\hat{\mu}$ (s.e.)	χ^2 (d.f.) (p-value)
Panel	511	-0.85	0.32	9.77 (12)
		(0.56)	(0.19)	(0.64)
Canada	158	-0.05	0.96	9.72 (2)
		(0.22)	(0.11)	(0.01)
Denmark	50	-1.18	0.81	2.16 (2)
		(1.44)	(0.25)	(0.34)
Finland	102	0.49	1.69	4.84 (2)
		(0.94)	(2.17)	(0.09)
France	82	-1.11	1.23	5.80 (2)
		(0.85)	(0.28)	(0.06)
Italy	113	-0.09	1.10	5.55 (1)
		(2.57)	(0.83)	(0.02)
Japan	75	2.91	0.59	4.08 (1)
		(5.13)	(0.75)	(0.04)
New Zealand	68	-1.00	1.03	2.96 (2)
		(0.53)	(0.08)	(0.23)
Sweden	74	-0.38	1.40	4.94 (1)
		(1.19)	(0.66)	(0.026)
UK	166	-0.22	1.15	7.48 (2)
		(0.33)	(0.26)	(0.02)

Notes: The United States is the reference country j in each case. The panel excludes Denmark and New Zealand, runs from 1981:II to 1999:III, and adopts a constant and lagged own-residuals as instruments. The instrument set for individual country estimates includes a vector of ones, the lagged growth rate of relative consumption, the lagged growth rate of relative government spending, and the lagged residual. The Italian, Japanese and Swedish cases exclude lagged relative consumption growth.

Table 2
Power Utility in Consumption

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{\alpha}$$

Country	T	$\hat{\alpha}$ (s.e.)	χ^2 (d.f.) (p-value)
Panel	511	-0.20 (0.51)	13.7 (13) (0.39)
Canada	158	0.01 (0.17)	8.69 (2) (0.13)
Denmark	50	-1.64 (1.52)	1.37 (1) (0.24)
Finland	102	0.04 (0.48)	6.04 (2) (0.05)
France	82	-0.21 (3.63)	5.36 (1) (0.02)
Italy	113	2.05 (0.65)	6.68 (2) (0.04)
Japan	75	1.79 (3.22)	2.71 (1) (0.10)
New Zealand	68	-1.03 (0.51)	3.04 (2) (0.22)
Sweden	74	-0.83 (1.33)	4.42 (1) (0.04)
UK	166	-0.39 (0.31)	7.44 (2) (0.02)

Notes: The United States is the reference country j in each case. The panel excludes Denmark and New Zealand, runs from 1981:II to 1999:III, and adopts a constant and lagged own-residuals as instruments. The instrument set for individual country estimates includes a vector of ones, the lagged growth rate of relative consumption, and the lagged residual. The Italy-US model substitutes lagged relative consumption growth with lagged relative employment growth.

Table 3
Power/Cobb-Douglas Utility in Consumption and Real Balances

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{1-\mu(1-\alpha)} \left(\frac{\Delta m_{it}}{\Delta m_{jt}}\right)^{-(1-\mu)(1-\alpha)}$$

Country	T	$\hat{\alpha}$ (s.e.)	$\hat{\mu}$ (s.e.)	χ^2 (d.f.) (p-value)
Panel	511	-1.18	0.87	13.54 (12)
		(0.51)	(0.11)	(0.33)
Canada	130	-0.09	1.00	10.25 (2)
		(0.26)	(0.20)	(0.01)
Denmark	50	-1.16	0.98	2.01 (2)
		(1.13)	(0.05)	(0.37)
Finland	94	-0.74	0.53	4.25 (1)
		(3.52)	(0.69)	(0.04)
France	82	-0.96	0.98	7.97 (2)
		(0.87)	(0.25)	(0.02)
Italy	93	-0.45	1.48	6.18 (2)
		(2.07)	(0.73)	(0.05)
Japan	75	1.21	-5.15	8.84 (2)
		(1.36)	(37.80)	(0.01)
New Zealand	68	-1.52	0.88	3.59 (2)
		(0.64)	(0.05)	(0.17)
Sweden	74	-0.77	1.03	4.58 (1)
		(1.25)	(0.22)	(0.03)
UK	150	-0.54	0.96	7.22 (2)
		(1.879)	(1.18)	(0.01)

Notes: The United States is the reference country j in each case. The panel excludes Denmark and New Zealand, runs from 1981:II to 1999:III, and adopts a constant and lagged own-residuals as instruments. The instrument set for individual country estimates includes a vector of ones, the lagged growth rate of relative consumption, the lagged growth rate of relative real balances, and the lagged residual. The Finnish and Swedish cases exclude lagged relative consumption growth.

Table 4
External Habit Persistence

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{\phi}\left(\frac{\Delta c_{it-1}}{\Delta c_{jt-1}}\right)^{-\theta}\kappa_{ij}$$

Country	T	$\hat{\phi}$ (s.e.)	$\hat{\theta}$ (s.e.)	$\hat{\kappa}$ (s.e.)	χ^2 (d.f.) (p-value)
Panel	504	1.84	1.52	1.00	6.36 (11)
		(0.44)	(0.40)	(0.01)	(0.85)
Canada	156	-0.06	0.07	1.00	7.57 (1)
		(0.19)	(0.15)	(0.01)	(0.01)
Denmark	48	-0.61	-0.65	1.01	1.82 (1)
		(1.45)	(1.53)	(0.02)	(0.18)
Finland	100	0.06	0.22	1.00	4.07 (1)
		(0.71)	(0.62)	(0.01)	(0.04)
France	80	-1.01	-0.67	1.00	5.77 (1)
		(0.92)	(0.86)	(0.01)	(0.02)
Italy	111	-1.18	0.37	0.99	4.14 (1)
		(1.16)	(1.10)	(0.01)	(0.04)
Japan	73	1.39	-1.76	1.01	2.53 (1)
		(1.35)	(1.39)	(0.01)	(0.12)
New Zealand	66	-1.12	0.12	1.00	5.99 (1)
		(1.15)	(1.26)	(0.01)	(0.01)
Sweden	72	-2.58	0.79	0.98	3.17 (1)
		(1.87)	(1.59)	(0.02)	(0.08)
UK	164	-1.08	0.89	0.99	4.84 (1)
		(0.55)	(0.46)	(0.01)	(0.03)

Notes: The United States is the reference country country j in each case. The panel excludes Denmark and New Zealand, runs from 1981:III to 1999:III, and adopts a constant and lagged own-residuals as instruments. The instrument set for individual country estimates consists of a constant, the lagged residual, and two lags of the growth rate in relative consumption.

Table 5
Endogenous Discount Rate

$$\Delta\left(\frac{e_{ijt}p_{jt}}{p_{it}}\right) = \left(\frac{\Delta c_{it}}{\Delta c_{jt}}\right)^{\alpha}\left[\frac{\Delta(1+c_{it-1})}{\Delta(1+c_{jt-1})}\right]^{\gamma}$$

Country	T	$\hat{\alpha}$ (s.e.)	$\hat{\gamma}$ (s.e.)	χ^2 (d.f.) (p-value)
Panel	504	1.53	-2.05	8.63 (12)
		(0.374)	(0.350)	(0.73)
Canada	157	-0.10	-1.28	8.36 (1)
		(0.22)	(0.95)	(0.00)
Denmark	49	0.26	-1.68	2.13 (1)
		(0.66)	(1.64)	(0.36)
Finland	101	-1.20	-0.54	2.93 (1)
		(2.64)	(1.54)	(0.09)
France	81	-0.32	-3.91	4.29 (1)
		(0.84)	(2.85)	(0.04)
Italy	112	9.94	-8.42	0.50 (1)
		(6.18)	(3.45)	(0.48)
Japan	74	-1.18	3.53	2.17 (1)
		(1.86)	(3.15)	(0.14)
New Zealand	67	-2.29	-0.32	2.13 (1)
		(1.06)	(0.62)	(0.14)
Sweden	73	-0.59	0.05	2.53 (1)
		(0.62)	(0.51)	(0.11)
UK	165	0.49	-0.59	3.87 (1)
		(0.83)	(0.65)	(0.05)

Notes: The United States is the reference country country j in each case. The panel excludes Denmark and New Zealand, runs from 1981:II to 1999:III, and adopts a constant and lagged own-residuals as instruments. The instrument set for individual country estimates consists of a vector of ones, the lagged residual, and the lagged growth rate in the real exchange rate.