Money and currency market implications of the long run risk models

Stefano d'Addona*

Abstract

Using postwar US and UK data, I study the implied interest rates and exchange rates in a simple long run risk (LRR) model with two countries. When a closed economy is considered, empirical estimates show that, as in standard consumption based models with power utility preferences, the movements of the implied risk free rate are entirely determined by the variations of the expected consumption growth. This leads to a negative relationship between LRR Euler equation rates and money market rates. Nevertheless, when the low frequency movements of consumption growth are explicitly included in the LRR estimates, the long run component of the consumption growth turns out to be a key element to capture the counter-cyclical variations of the money market rates. Turning to an open economy setting, I find supporting evidence in favor of the LRR model. Again, the long run component of the consumption growth plays a determinant role in accounting for foreign exchange market movements.

JEL: E44, F31, F41, G12

*Department of International Studies, University of Rome 3, Via G. Chiabrera, 199, I-00145 Rome; +39-06-5733-5331; daddona@uniroma3.it

1 Introduction

The consumption Euler equation has become a central paradigm for several macroeconomic models both in the asset pricing and the real business cycle literature. In these class of models, the stochastic discount factor provides a core relation between the money market rate and the implied riskless rate. It follows that the growth expectations of marginal utility in the pricing kernel should reflect the stance of monetary policy so that the implied risk free return should match the observed level of money market rates. Moreover, in a two-country setting, if markets are complete, the ratio of the stochastic discount factors for domestic and foreign currency denominated assets should reflect the currency depreciation rate (see Backus et al., 2001).

This work is concerned with the monetary and exchange market implications of a simple long run risk (LRR) model (cf. Bansal and Yaron, 2004). As further discussed below, if agents have a preference for early resolution of uncertainty and are sufficiently risk adverse, long run risk helps explain the equity premium in consumption based asset pricing models using a persistent component in the process for the aggregate consumption growth. As in any standard macroeconomic model, the Euler equation implied by a LRR model provides a direct implication for monetary policies and foreign exchange markets in an open economy setting. By using postwar US and UK data, this link is used to empirically test the capability of the model to capture the historical behavior of the interest rates in the two countries and their exchange rate.

Evidence of the empirical weakness of the consumption-based Euler equation is not new to the asset pricing literature. The "equity premium puzzle" documents the mismatch between the sizeable excess return observed in the US market and that generated by the standard consumption CAPM (See Mehra and Prescott, 1985). The "risk free rate puzzle" (see Weil, 1989) highlights the difference between the unconditional mean of the risk free return implied by the consumption-based Euler equation and the observed Treasury bill rate.

Changing the preference structure or the economic environment is useful in reconciling model-generated returns with the unconditional moments of financial data: habit formation has contributed to the resolution of the equity premium puzzle quite successfully (see Abel, 1990, Campbell and Cochrane, 1999). Further, Tallarini (2000) shows that introducing recursive preferences, such as Epstein and Zin (1989), in a real business cycle model improves the model's asset market predictions.

Nonetheless several studies warn about the empirical inconsistencies which threaten these assumptions. Using US data, Canzoneri et al. (2007) find that the riskless return implied by the Euler equations of the above cited specifications and the observed pattern of the federal funds rate do not coincide. Also, the unconditional moments of implied money market rates exceed, in most cases, their empirical counterpart and this evidence holds true whether the stochastic discount factor is nominal or real[1]. As far as foreign exchange markets are concerned, Backus and Smith (1993) show evidence of poor correlation detected among exchange rates and consumption differentials and argue that this finding creates a gap between international prices and quantities, at odds with standard power utility models. Brandt et al. (2006) find that, due to the low cross-country co-movement of short term consumption streams, the difference in the stochastic discount factors is too volatile to capture the smoothness of the US dollar depreciation rate.

Abstracting from the preference structure, the core of these inconsistencies stems from the interaction between macroeconomic variables and the interest rates implied by the consumption Euler equation. Atkeson and Kehoe (2009) suggest that the Euler equation misses the real link between the policy instrument and the economic forces which drive the monetary policy. In particular, standard asset pricing models relate the risk free returns to the short term fluctuations of the expected consumption growth. In this context, risk plays no role and the conditional variance of

[1] Ahmad (2005) confirms that this average spread is not an isolated artifact of the US and that it concerns the other six G7 countries as well.

consumption, which enters the log of the Euler equation, is constant.

A historical analysis of the US money market rates reveals instead that the response of the Federal Reserve Bank to the expected changes of consumption growth is smooth, whereas the monetary policy mainly reacts to the business cycle fluctuations in real risk. Thus, to reflect the stance of monetary policy, the Euler equation should be developed in a framework where the risk free returns react smoothly to the variations of the expected consumption growth and the volatility of consumption varies at a business cycle frequency.

The long run risk model of Bansal and Yaron (2004) seems to meet these requirements. Differently from the standard asset pricing models, Bansal and Yaron (2004) specify consumption as containing a small and persistent component that accounts for the low frequency movements of the economic growth. The conditional volatility of consumption is allowed to vary over the business cycle, introducing fluctuating economic uncertainty and time-varying risk. In this framework, the Epstein and Zin (1989) preferences emphasize the role of long run growth and economic uncertainty as additional sources of risk, and smooth the response of the risk free asset to the short term consumption variations.[2]

My empirical analysis shows indeed that the long run characteristics of the LRR model contribute to the macroeconomic consistency of the model-generated risk free rate. In particular, when the low frequency movements of consumption growth are explicitly included in the Euler equation, the correlation between the real risk free return and the the money market rates gets positive and the two series behave similarly. Turning to an open economy setting, evidence is provided that the same LRR specification accounts for foreign exchange market movements if the focus of

[2]Colacito and Croce (2011) introduce the LRR model in a two country setup with complete markets. The authors show that, when long-run consumption is highly persistent and long-run risks are assumed to be strongly correlated, the model is able to generate co-moving stochastic discount factors and exchange rate fluctuations close to the unconditional moments observed in the data. More recently, Bansal and Shaliastovich (2007) show that introducing stochastic volatility in the two-country setting of Colacito and Croce (2011) helps explain a wide array of asset pricing puzzles while accounting for exchange rate volatility.

the analysis is on the last two decades of data.

The remainder of the paper is organized as follows: Section 2 will specify the consumption Euler equation for the standard power utility case, the LRR model of Bansal and Yaron (2004), and the two-country model of Colacito and Croce (2011); Section 3 explains in detail the estimation procedure and describes the data; Section 4 presents the results; Section 5 concludes.

2 Riskless interest rates and foreign exchange market rates

Let us consider an endowment economy with two countries. As the two economies are characterized as symmetric ones, preferences, endowments, and returns for the home and the foreign country have the same system of equations indexed by f for the latter.

A representative household maximizes the following recursive utility function

$$V_t = \left[(1-\beta)C_t^{1-\frac{1}{\psi}} + \beta\mathcal{R}_t(V_{t+1})^{1-\frac{1}{\psi}}\right]^{\frac{1}{1-\frac{1}{\psi}}}. \tag{1}$$

The operator $\mathcal{R}_t$ makes a risk adjustment to the date $t+1$ continuation value. The risk adjustment is given by

$$\mathcal{R}_t(V_{t+1}) = \left(\mathrm{E}_t\left[V_{t+1}^{1-\gamma}\right]\right)^{\frac{1}{1-\gamma}}.$$

The parameter γ is the coefficient of risk aversion (RA), β denotes the household's subjective discount factor while the elasticity of intertemporal substitution (EIS) is given by ψ.

The representative household is assumed to allocate her disposable income between consumption and two one-period bonds: specifically, one bond that pays out one consumption unit (the real asset), the other that returns one unit of country's currency (the nominal asset).

2.1 The power utility case

The function (1) reduces to a monotone transformation of the standard power utility function for $\psi = \gamma^{-1}$. According to the standard power utility preference structure, consumption utility is defined as $u(c_t) = \frac{C_t^{1-\gamma}}{1-\gamma}$, where the marginal utility equals $u'(C_t) = C_t^{-\gamma}$.

Conforming to the first order condition of the household's maximization problem in (1), the gross real risk free return $R_t^* = 1 + r_t^*$ satisfies

$$R_t^* = \beta^{-1} E_t \left[\left(\frac{C_{t+1}}{C_t} \right) \right]^{\gamma} \tag{2}$$

where the right-hand side of equation (2) is the conditional expectation of the inverse of the stochastic discount factor $M_{t+1} = \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma}$.

While the nominal rates, defined as $1 + i_t^* = (1 + i_{t+1}^*) * (1 + \pi_{t+1})$, where i_t^* is the net nominal risk free return, $1 + \pi_{t+1} = (P_{t+1}/P_t)$ is the inflation rate, and P_{t+1} is the price level, satisfies

$$1 + i_t^* = \beta^{-1} E_t \left[\left(\frac{C_{t+1}}{C_t} \right)^{\gamma} \left(\frac{P_t}{P_{t+1}} \right)^{-1} \right]. \tag{3}$$

If consumption and inflation are assumed to be jointly log-normal, the log-linearized Euler equations in (2) and (3) take the form

$$r_t^* = -\ln \beta + \gamma E_t[\Delta(c_{t+1})] - \frac{\gamma^2}{2} \sigma_t^2 [\Delta(c_{t+1})] \tag{4}$$

and

$$\begin{aligned} i_t^* = & -\ln \beta + \gamma E_t[\Delta(c_{t+1})] + E_t[\pi_{t+1}] - \frac{\gamma^2}{2} \sigma_t^2 [\Delta(c_{t+1})] \\ & - \frac{1}{2} \sigma_t^2 [\pi_{t+1}] - \gamma cov_t [c_{t+1}, \pi_{t+1}] \end{aligned} \tag{5}$$

where $[\Delta(c_{t+1})]$ is the (log) consumption growth rate, $\sigma_t^2[\Delta(c_{t+1})]$ and $\sigma_t^2(\pi_{t+1})$ denote the conditional variances of the consumption growth and of the inflation rate, and the term $cov_t[c_{t+1}, \pi_{t+1}]$ captures their conditional covariance.

2.2 The long run risk model

Abstracting from the special case $\psi = \gamma^{-1}$, the derivation of the real risk free return proceeds from the log-linearization of the following Euler equation (see Epstein and Zin, 1989)

$$R_t^* = E_t\left[\beta^\theta \left(\frac{C_{t+1}}{C_t}\right)^{-\frac{\theta}{\psi}} R_{a,t+1}^{-(1-\theta)}\right]^{-1} \tag{6}$$

from which the log of the stochastic discount factor is

$$m_{t+1} = \theta \ln \beta - \frac{\theta}{\psi}\Delta c_{t+1} + (\theta - 1)(r_{a,t+1}) \tag{7}$$

where $\theta = \frac{1-\gamma}{1-1/\psi}$ and $r_{a,t+1}$ is the log of the *unobservable* gross return on the aggregate consumption claim.

Bansal and Yaron (2004) specify consumption growth according to the following exogenous law of motion

$$\begin{aligned}
\Delta[c_{t+1}] &= \mu_c + x_t + \sigma_t \eta_{t+1} && (8)\\
x_{t+1} &= \rho_x x_t + \varphi_e \sigma_t e_{t+1} && (9)\\
\sigma_{t+1}^2 &= \sigma^2 + v_1(\sigma_t^2 - \sigma^2) + \sigma_w w_{t+1} && (10)\\
&\quad \eta_{t+1}, e_{t+1}, w_{t+1} \sim i.i.d.N(0,1)
\end{aligned}$$

The conditional expectation of the consumption growth (that is, $\mu_c + x_t$) results from the combination of the unconditional average, μ_c, and a slow-moving predictable component, x_t. The state variable x_t characterizes the long run properties of the consumption growth process, with the parameter ρ_x measuring its persistence. While the shock η_{t+1} represents a standard high frequency innovation in short term consumption, the innovation term e_{t+1} captures the long run risk of consumption prospects, and φ_e determines its predictability. To account for the economic uncertainty affecting consumption growth, its variance σ^2_{t+1} is defined as an AR(1) process and a persistence parameter v_1. The error term w_{t+1} is a shock to the economic uncertainty and σ_w captures its unconditional volatility.

Bansal and Yaron (2004) show that the real risk free return r^*_t satisfies:

$$r^*_t = -\theta \log(\beta) + \frac{\theta}{\psi} E_t[\Delta c_{t+1}] + (1-\theta) E_t[r_{a,t+1}] + \frac{1}{2} Var_t \left[\frac{\theta}{\psi} \Delta g_{t+1} + (1-\theta) r_{a,t+1} \right] \tag{11}$$

It is worth noting that, besides the consumption dynamics above, the complete characterization of the LRR stochastic discount factor relies on i) an estimable solution for $r_{a,t+1}$, ii) an estimation of the variance term in equation (11). The return $r_{a,t+1}$ is derived by the approximation:

$$r_{a,t+1} = k_0 + k_c z_{t+1} - z_t + \Delta c_{t+1} \tag{12}$$

which is a function of the log price-to-consumption ratio z_t with approximating constants, $k_c = \exp(\tilde{z})\ 1 + \exp(\tilde{z})$ and $k_0 = \log(1 + \exp(\tilde{z})) - k_c \tilde{z}$ (see Campbell and Shiller, 1988).

Bansal and Yaron (2004) show that the variance term $Var_t \left[\frac{\theta}{\psi} \Delta g_{t+1} + (1-\theta) r_{a,t+1} \right]$ is equal to the conditional variance of the stochastic discount factor which links the riskless return to the economic structure and the market compensation for consumption risks. That is

$$Var_t[m_{t+1}] = (\lambda^2_{m,\eta} + \lambda^2_{m,e})\sigma^2_t + \lambda^2_{m,w}\sigma^2_w \tag{13}$$

where $\lambda_{m,\eta}$,$\lambda_{m,e}$, and $\lambda_{m,w}$ are respectively the unit prices for the short-run risk, the long-run risk, and the volatility risk. Expressions for the market prices of risk are derived in terms of the parameters of the consumption dynamics and the preference structure. Namely $\lambda_{m,\eta} = \gamma$, $\lambda_{m,e} = \left(\gamma - \frac{1}{\psi}\right)\left[\frac{k_c \varphi_e}{1-k_c\rho}\right]$, and $\lambda_{m,w} = \left(\gamma - \frac{1}{\psi}\right)(1-\gamma)\left[k_c\left(1+\frac{k_c\varphi_e}{1-k_c\rho}\right)^2 \; 2(1-k_c v_1)\right]$.

2.3 The foreign exchange market

It is well known (see Backus et al., 2001) that if markets are complete, the exchange rate growth would equal the ratio of the stochastic discount factors for foreign currency and domestic currency denominated assets. Correspondingly, the growth of the log-exchange rate, Δe_{t+1}, would satisfy

$$\Delta e_{t+1} = m^f_{t+1} - m^h_{t+1}, \tag{14}$$

where the superscripts h and f indicate the home and the foreign country respectively. In the special case of power utility functions, the dynamics of the stochastic discount factor are solely determined by short term consumption prospects. In accordance with this condition, the currency valuation results from the cross-country difference of current consumption growth rates

$$\Delta e_{t+1} \propto \Delta(c^f_t) - \Delta(c^f_t)$$

As in Colacito and Croce (2011), I add to this baseline specification a long run risk component. In each of the two countries considered, domestic consumption dynamics are defined according to the following two-country environment:

$$\Delta[c^i_{t+1}] = \mu^i_c + x^i_t + \varepsilon^i_{c,t+1} \tag{15}$$

$$x^i_{t+1} = \rho^i_x x^i_t + \varepsilon^i_{x,t+1} \tag{16}$$

$$\xi \sim i.i.d.N(0,\Sigma)$$

where the vector $\xi = [\varepsilon^h_{c,t+1} \quad \varepsilon^h_{x,t+1} \quad \varepsilon^f_{c,t+1} \quad \varepsilon^f_{x,t+1}]$ collects the short and long risks affecting the home and foreign country.

Σ is the variance-covariance matrix of the shocks to the economy which are allowed to be cross-country correlated. Specifically volatilities are indicated by σ_c^i and σ_x^i respectively, while ρ_c^{hf} captures the cross-country correlation between the short-term innovations in consumption growth ($\varepsilon_{c,t+1}$ and $\varepsilon_{c,t+1}^f$), and ρ_x^{hf} denotes the cross-country correlation of long-run risks ($\varepsilon_{x,t+1}$ and $\varepsilon_{x,t+1}^f$).

From the first-order linear approximation of the model (see Colacito and Croce, 2011), the log of the country-specific stochastic discount factor is derived as:

$$m_{t+1} = log\delta - \frac{1}{\psi}x_t + k_c\frac{1-\gamma\psi}{\psi(1-\rho_x k_c)}\varepsilon_{x,t+1} - \gamma\varepsilon_{c,t+1} \tag{17}$$

where k_c is the approximation constant introduced in equation (12), and the superscript i is suppressed for the sake of readability. Accordingly, the exchange rate growth is expressed in terms of the relative valuation of both current and long-term consumption streams.

3 Estimation and data

As a first step in the empirical examination, I estimate the riskless returns implied by the Consumption CAPM with standard power utility preferences and relate them to the observed money market rates. This preliminary check allows us to show further evidence of the mismatch between the two interest rates series documented by the previous research on US data[3]. In particular, we expect the proportional link inferred by the model between risk free returns and short-term consumption growth to be insufficient for capturing the cyclical stance of monetary policy.

Later, we turn to the long run risk model (LRR) of Bansal and Yaron (2004) and its subsequent modifications (Bansal et al., 2007). The purpose is to investigate whether i) relaxing the restriction $\gamma = \psi^{-1}$; ii) introducing an explicit formulation

[3]See Canzoneri et al. (2007), Ahmad (2005) and Reynard and Schabert (2010) for some quantitative estimates.

for long-term consumption growth into the pricing kernel; and iii) accounting for economic uncertainty, reconciles the model implied risk free rate with the data.

Finally, to complete the empirical investigation, we analyze the international implications of both models applied to foreign exchange markets.

3.1 Estimation procedure: the power utility case

Following the methodology of Ahmad (2005), the first and second moments in equations (4) and (5) are derived from the vector autoregressive (VAR) model

$$Z_t = A_0 + AZ_{t-1} + \varepsilon_t \qquad \varepsilon_t \sim N(0, \Omega) \tag{18}$$

that is assumed to describe the exogenous dynamics of consumption and inflation. Specifically, Ω denotes the variance-covariance matrix, and Z_t includes the real log consumption c_t, the inflation rate π_t, the money market rate, the log of real disposable income per capita, a monetary aggregate, and the lags of all these variables.

The conditional expectation terms are obtained from the OLS one-step-ahead linear predictions of the first two equations in the VAR (18). The conditional second moments are retrieved from the variance-covariance matrix according to the following array

$$\sigma_t^2 c_{t+1} = e_1' \Omega e_1' \tag{19}$$

$$\sigma_t^2 \pi_{t+1} = e_2' \Omega e_2' \tag{20}$$

$$Cov_t[c_{t+1}, \pi_{t+1}] = e_1' \Omega e_2' \tag{21}$$

where the indicator vectors e_1' and e_2' pick up only the first and second elements of its diagonal[4].

[4]It is worth noting that this estimation set-up implies that the conditional variance of the stochastic discount factor is constant.

The risk free returns resulting from the standard CCAPM in (4) and (5) are finally compared to national money market rates data. In particular, the net nominal risk free rate i_t^* is compared to the nominal annualized money market rate i_t, while the net real risk free return r_t^* is compared to the *ex ante* real money market rate. The latter is constructed using the conditional expected inflation predicted by equation (18).

Finally, in line with previous studies (see Canzoneri et al., 2007), the coefficient of relative risk aversion γ is set equal to 2 and the subjective discount factor β equals 0.993.

3.2 Estimation procedure: the long run risk model

To fully assess the effects of long run consumption prospects and the economic uncertainty on the behavior of the money market rates, we estimate three concurrent specifications of the LRR model.

1. **Baseline calibration**

   According to the original specification introduced by Bansal and Yaron (2004), we firstly proceed with the baseline calibration of the equations (8)-(9)-(10), the choice of the preference parameters, and the estimation of the time-varying consumption variance, σ_t^2.

   Table 1 reports the parameters which describe the dynamics of consumption and the investor's preferences. The model is calibrated on a monthly basis in order to fit the monthly decision interval of the representative investor and avoid model mis-specification[5]. The parameters are set to simultaneously match the salient features of consumption and asset pricing dynamics in US and in UK data[6].

[5] Bansal et al. (2007) warn that the misalignment between the sampling frequency of consumption data and the agent's decision interval may lead to substantial biases on the economic plausibility of the model and the interpretation of the results.

[6] For a detailed description of the calibration procedure we applied, see Bansal and Yaron (2004).

[Table 1 about here.]

The choice of the preference parameters plays a key role in the long-term configuration of the model. In fact, a well know feature of the Epstein and Zin (1989) preferences is that the magnitude of the RA relative to the EIS governs the representative agent's timing of the uncertainty resolution.

Bansal and Yaron (2004) assume that the representative household prefers an *early* resolution of the uncertainty ($\gamma > \frac{1}{\psi}$) with $\psi > 1$. Under this condition, the representative agent gives higher weight to those consumption risks perceived as more long-lasting. Conforming to this configuration, we closely follow the original calibration of the Bansal and Yaron (2004) model and set the coefficient of risk aversion to 10 and let the EIS parameter be 1.5. The subjective discount factor is set to 0.998.

To complete the estimation of the model, the time series of the state variables σ_t^2 and x_t are needed. The variance σ_t^2 is predicted using a GARCH(1,1) for the autoregressive process of the consumption growth. Finally, the full characterization of the long run consumption growth relies entirely on the calibration of the parameters ρ_x and ϕ_e, which result from its autoregressive process.

2. Recovering long-run consumption growth

More recently, Bansal et al. (2007) have shown that the low frequency component x_t can be directly extracted from consumption and financial data. Hence, to provide a more concrete measure of long run consumption, we decided to add to the baseline calibration of the model and recover x_t from the data. Following Bansal et al. (2007), I regress consumption growth on the ex-ante real risk free rate and the price-dividend ratio[7]. That is I perform

$$\Delta c_{t+1} = \zeta Y_t + \sigma_t \eta_{t+1} \tag{22}$$

and take the expectation of model predictions

$$x_t = \zeta Y_t \tag{23}$$

where $Y_t = \begin{bmatrix} 1 & z_{pd,t} & r_t^f \end{bmatrix}'$, $z_{pd,t}$ is the price-dividend ratio, and r_t^f is the observed real risk free rate. It follows that the conditional expected consumption growth can be explicitly computed as the sum of x_t and μ_c.

Again the parameters ρ_x and φ_e are set according to the first order autoregression of x_t. At a monthly frequency, ρ_x is equal to 0.799 and φ_e is 0.083, suggesting that the low frequency component of the consumption growth is less persistent than the baseline LRR calibration and slightly more predictable (cf. table 1).

3. **Abstracting from the economic uncertainty**

   Finally, to assess the interaction of the economic uncertainty with the empirical performance of the LRR model, we estimate the risk free rate in (11) abstracting from the time variations of the consumption volatility. This step requires a brief reconfiguration of the equations above. Specifically, the conditional variance σ_t^2 is substituted by its unconditional mean σ^2. The purpose is to check whether the results are somehow affected by the heteroschedasticity of consumption growth.

[7] To obtain the ex-ante real risk free rate r_t^f, I followed Bansal et al. (2012) and regressed the ex-post real short term rate on its nominal counterpart and the past annual inflation. The ex-post real short term rate is given by the difference between the nominal rate and the inflation rate.

3.3 Estimation procedure: Exchange rates

As discussed in section 2.3, the real effective exchange rate implied by the LRR model results from the first order approximation of the stochastic discount factor reported by equation (17).

The model is empirically tested by estimating the relevant parameters in the system of equations (15)-(16)-(17) with a Kalman filter procedure with the observation equation and state equation specified as follows:

$$\begin{aligned} C_t &= A + \Phi X_t + R\epsilon_{c,t}, \qquad (24) \\ X_{t+1} &= FX_t + Q\epsilon_{x,t}, \end{aligned}$$

Where $C_t = \begin{bmatrix} \Delta[c_t^h], & \Delta[c_t^f] \end{bmatrix}'$ is the vector of observed variables at time t, $X_t = \begin{bmatrix} x_t^h, & x_t^f \end{bmatrix}'$ is the vector of unobservable state variables, A is a column vector containing μ_c^h and μ_c^f, Φ is a 2×2 identity matrix, F is a 2×2 diagonal matrix containing the autoregressive parameters for the two unobservable processes, R and Q are the two variance-covariance matrices, and $\epsilon_{c,t}$ and $\epsilon_{x,t}$ are white noise bivariate independent vectors. Table 1, Panel B, reports the shared set of preference and consumption parameters. Specifically, consumption dynamics are calibrated letting the two economies be asymmetric while preference parameters are assumed to be the same in both countries. As a term of comparison we also estimated a "restricted" version of the model where the cross-country correlation in short-term and long-term risks are fixed as in Colacito and Croce (2011).

3.4 Data

Estimations are based on United States and United Kingdom postwar data. Specifically, the US is denoted as the home country and UK as the foreign one. When money market rates are concerned, we treat them as closed economies. When exchange rates are estimated, I define the United States and the United Kingdom as open and symmetric economies. In order to maintain continuity with the existing literature, US estimates will cover the overall quarterly period 1964q1-2011q4[8]. Due to data availability, I restricted the observation set of the United Kingdom to the time span 1981q1-2011q4, instead.

Consumption and disposable income. All data for the US consumption on non-durable goods and services and the real disposable income is acquired from the NIPA (*National Income and Product Accounts*) Tables of the *Bureau of Economic Analysis* (BEA). Per capita figures are worked out using population data from the *U.S. Census Bureau.* I obtained the nominal and real consumption series and the data for the UK real disposable income from the UK's Office for National Statistic (ONS) *Quarterly National Accounts.* UK population figures are taken from *Eurostat.*

Consumption-based inflation. Inflation is computed using a consumption-based price index. This index results from the ratio between the nominal and real consumer spending components above. The inflation rate is obtained as the change in the price level over the previous quarter.

Interest rates and the money stock. US data for the money market rate (i.e, the federal funds rate), the 3-month Treasury bill rate and the monetary aggregate M2 are extracted from the *Federal Reserve Economic Data* (FRED) of the St. Louis Fed. As a measure for the UK money market rate we use the LIBOR (London Interbank Offered Rate), acquired from *Eurostat.* The UK 3-month Treasury bill rate is provided by the *Bank of England Database*, the money stock M4 by the ONS's *Financial Statistics.*

Stock prices, dividends, and the exchange rate. Us data for the S&P500 Composite index and dividends is available on the Robert Shiller's *U.S. Stock Price Database.* UK stock prices (specifically, the FTSE 100 index) are provided by the *Financial Times* while dividends are obtained from the ONS. The real dollar-pound

[8] Ahmad (2005) considers the quarterly time period 1964Q1-2000Q4; Canzoneri et al. (2007) extend it to 2002Q4. To get comparable results, the full sample starts in 1964Q1 as well and drops in 2011Q4, including more recent observations.

exchange rate is obtained by multiplying the nominal dollar-per-sterling rate by a price deflator. The price deflator is constructed as the ratio of the UK and the United States consumer price index[9]. The nominal exchange rate is provided by the FRED database.

4 Results

4.1 Riskless returns and monetary policy stance: the power utility case

Turning to the results, we firstly examine the riskless returns produced by the consumption CAPM when agents have power utility preferences.

[Table 2 about here.]

As expected, as far as the US economy is concerned, the model does not succeed in fitting the data. The summary statistics in the upper panel of Table 2 signal that the *ex-ante* real federal funds rate, r_t, and the risk free rate implied by the model, r_t^*, diverge: although the model matches the overall volatility of the data, the unconditional mean of the risk free return exceeds the observed average of the real federal funds rate by 3%. The last row of the US panel in Table 2 reports the correlation coefficient between the model implied rate and the federal funds rate, confirming their different behavior: The two interest rate series are nearly uncorrelated with a correlation equal to -0.02 that is statistically non different from zero[10].

By graphical inspection (cf. Figure 1) the time series comparison on the last decade reveals that this average gap rises between 2001 and 2011, when the real risk

[9] The UK consumer price index is available on the OECD's *Main Economic Indicators*; as a measure for the US price level, we use the *All Urban Consumers price index*, published by the *Bureau of Labor Statistics*

[10] These results are robust if the model is checked as the 3-month Treasury bill rate is assumed to be the appropriate counterpart for the risk free rate.

free rate implied by the model averages 4.5% while the ex-ante real federal funds rate turns out to be negative.

During the two most recent recessions[11], the sharp fall registered in the federal funds rate systematically corresponds to opposite increases in the model implied rates. This regularity implies that some causal linkage may exist between the magnitude of this difference and the stance of the monetary policy.

[Figure 1 about here.]

Analyzing the time variations of the spread between the data and the model-generated returns supports this intuition. Table 3 provides the output obtained from regressing the real spread $(r_t^* - r_t)$ on its four lags and the real federal funds rate. The coefficient of fed funds turns out to be negative ($\beta_{r_t} = -1.45$) and strongly significant. This suggests that, when a monetary policy easing is managed to support the recovery of the economic system, the risk free return obtained from the Euler equation (4) directs toward an opposite path.

[Table 3 about here.]

Estimates based on nominal rates are reported on the last two columns of Table 2 (upper panel). Again, the mean of the observed nominal realizations of the fed funds rate and the mean of the model implied nominal rate i_t^* diverge with the latter being, on average, 4% greater.

[Figure 2 about here.]

Interestingly, the nominal series are positively correlated ($\rho = 0.44$). According to Reynard and Schabert (2010), this positive co-movement may be due to the latent influence of inflation trends leading the two nominal interest rates in the same direction. If this were the case, the positive correlation between i_t^* and i_t^*

[11] Recession dates are from the National Bureau of economic research (NBER)

would just be a misleading signal of the effective connection between the data and model-generated returns. A natural way to assess this issue is by inspecting the trend and cycle properties of the series. Thus the trend and cyclical components of the three variables (inflation rate, implied nominal rate, and nominal market rate) were extracted with a Hodrick and Prescott (1997) filter[12]. A comparison of the inflation trend with the trends of the money market rate and of the implied nominal risk free rate supports the above intuition. The resulting correlation coefficients ($\rho = 0.74$ and $\rho = 0.70$ respectively) confirm that the underlying dynamics of the two nominal interest rates are equally affected by the trend in price movements. Moreover, the cyclical components of i_t^* and i_t^* follow a different behavior: the correlation coefficient turns negative ($\rho = -0.22$) and statistically different from zero. Finally, regressing the nominal spread ($i_t^* - i_t$) on its four lags and the nominal federal funds rate produces similar results to those obtained for the analysis of real returns. Once again the coefficient on the nominal federal funds rate, β_{i_t}, is negative and significant ($\beta_{i_t} = -1.52$).

Results on the UK economy are reported in the lower panel of Table 2. The unconditional mean of the implied real rate exceeds the mean of the real LIBOR by 2%. Further, it is more volatile than the observed money market rate. Comparing the nominal risk free return and the nominal LIBOR yields similar results. The correlation between the two real series turns out to be positive ($\rho = 0.33$). However, this is not robust to the sample considered. In fact, when the sample is restricted to the last decade of data (2001-2011) the correlation coefficient drops to $\rho = 0.17$ and is not significantly distinguishable from zero.

[Figure 3 about here.]

The nominal estimates confirm the results obtained on the US economy. The correlation between the nominal implied rate and the nominal LIBOR appears pos-

[12]Given the quarterly frequency of the observations, the smoothing parameter is set equal to 1600.

itive (around $\rho = 0.50$). Nonetheless, Figure 4 suggests that the two interest series cyclically diverge. As a robustness check, I applied a Hodrick and Prescott (1997) filter to i_t^* and i_t and compared their cyclical components. The resulting correlation coefficient turns negative ($\rho = -0.38$) and statistically different from zero at a 1% significance level. Also when the trend components are compared with the trend extracted from the inflation rate, the resulting correlation coefficients are high and positive ($\rho = 0.49$ and $\rho = 0.73$ respectively).

Instead, the analysis of the relation between the model-data spread and the stance of monetary policy leads to different conclusions for the UK data. As shown by the rightmost panel of Table 3, a regression between the real time-varying spread $(r_t^* - r_t)$ on its four lags and the *ex-ante* real LIBOR suggests that there is no regularity between the difference of the two interest rates and the stance of monetary policy: the regression coefficient β_{r_t} is negative but not significant ($\beta_{r_t} = -0.26$).

This is confirmed by the output of the nominal spread regression. Results (see Table 3) indicate that the gap between the model rate and the data is not affected by the money market rate level (the coefficient $\beta_{r_t} = 0.01$ and is not statistically significant).

[Figure 4 about here.]

4.2 Riskless returns and money market rates: the long run risk model

Table 4 (column 2) shows the summary statistics resulting from the baseline calibration of the long run risk model. Compared to the power utility case, the LRR estimates reduce the average difference between the model returns and the money market rates. However, the model still does not succeed in fitting the data as the interest rate series still diverge. On the one hand, the US risk free return is on average twice the federal funds rate (3,31% compared to 1,88%) and less volatile; on the other, the unconditional mean of the UK implied rate sharply falls to 2.76.

Figures 5 and 6 plot the two interest rates.

Focusing on the US, the short term movements of the risk free rate bear no resemblance to the pattern of the federal funds rate. The correlation coefficient, $\rho = -0.0261$ confirms that the two time series are almost uncorrelated. Similarly, the correlation between the UK implied rate and the real LIBOR approaches zero ($\rho = 0.01$).

[Table 4 about here.]

[Figure 5 about here.]

[Figure 6 about here.]

Equation (11) implies that the movements of the risk free rate are governed by the long run fluctuations of the expected consumption growth. This is a prominent difference compared to the power utility case where just the high frequency shocks to consumption are considered. Yet, the results show that the short term variations of consumption still play a dominant role for the determination of the risk free return and suggest that the baseline calibration of the LRR model does not capture the long run consumption dynamics.

Interestingly, the third column of Table 4 shows that when the long run component x_t is explicitly included in the model estimation the results sharply change. The US and the UK correlation coefficients turns positive (0.43 and 0.20). Figures 7 and 8 show that the model implied risk free rates closely follow the path of their empirical counterparts. These results suggest that the long run characteristics of the expected consumption growth are crucial for the macroeconomic consistence of the risk free estimates.

Nevertheless, as far as the US economy is concerned, the unconditional moments of the two interest rates still diverge. The mean of r_t^* rises up to 4.20, so that the difference between the average levels is higher than previous estimates. Meanwhile, the standard deviation of r_t^* lowers to 1.27.

In contrast, the UK implied rate matches the data almost exactly, with its unconditional mean approaching 4.76%(compared to 4.08%) and its standard deviation reducing to 2.434 (compared to 2.362 in the data).

[Figure 7 about here.]

[Figure 8 about here.]

While the above results suggest that the state variable x_t contributes to the macroeconomic consistence of the implied risk free rate, the outcome may as well be influenced by the heteroschedasticity modeled in the volatility of the consumption growth (σ_t^2). This may be at odds with the previous literature: according to Beeler and Campbell (2009), the effects of the economic uncertainty on the consumption-saving decisions of the representative agent should be not significant.

To explore the interaction between the long run growth and the economic uncertainty in the determination of the risk free rate, I estimate the LRR model abstracting from the time variations of the consumption volatility. Column 4 of Table 4 reports the estimates. Results on US economy are not affected by this modification. The mean of the risk free rate lowers to 3.87 with a standard deviation is 1.28. The correlation is slightly higher ($\rho = 0{,}44$). Also UK estimates are unchanged.

Summarizing, as for the power utility case, the movements of the risk free rate implied by LRR models are entirely determined by the variations of the expected consumption growth. This result seems to be at odds with the conclusions of Atkeson and Kehoe (2009) who state that time-varying risk is a central indicator for the monetary policy of the Federal Reserve. Nevertheless, differently from the standard power utility specification, the long run component of consumption growth is a key element to capture the counter-cyclical variations of the money market rates.

[Table 5 about here.]

4.3 Exchange rates

Table 6 reports the estimates of the real exchange rate growth implied by the equation (14) for the power utility case and for two specifications of the LRR model. In the vein of Colacito and Croce (2011) the restricted model implies complete symmetry between the two economies and fixes ρ_x and ρ_c to 1 and 0.3 respectively. The full model instead does not impose any restriction on the estimated parameters.

When the whole sample is considered, the three models display a low correlation with the observed time series of the real USD/GBP exchange rate. These results overwhelmingly lead to the conclusion that all models fail to mimic the exchange rate market. When data before the first quarter of 1990 are cut off from the sample, the correlation between the full LRR model and the data is positive and statistically distinguishable from zero. This is consistent with the sharp increase in international flows, both on financial markets and commodities, documented in the literature (e.g. Baier and Bergstrand, 2007 and Razin and Sadka, 2007).

[Table 6 about here.]

[Figure 9 about here.]

[Figure 10 about here.]

[Figure 11 about here.]

5 Concluding remarks

I studied the macroeconomic implications of a class of consumption based asset pricing models where a persistent component in the process for the aggregate consumption growth is accounted for. Using postwar US and UK data I estimated the implied interest rates in the long run risk model introduced by Bansal and Yaron (2004) and its subsequent modifications (Bansal et al., 2007). Then I extended the estimates to the exchange rate market, in a two countries setting.

When the closed economy money market is considered, I show that, as in the standard consumption based CAPM, the movements of the risk free rate implied by a baseline LRR model are entirely determined by the variations of the expected consumption growth. Nevertheless, differently from the standard CCAPM, the long run characteristics of the LRR model contribute to the macroeconomic consistence of the model-generated risk free return. In particular, when the low frequency movements of consumption growth are explicitly included in the Euler equation, the correlation between the implied real risk free rate and the federal funds rate gets positive and the two interest rate series move in the same direction. Turning to an open economy setting, I find evidence that the LRR model accounts for foreign exchange market movements when the last two decades of data are considered.

References

Abel, Andrew B., "Asset Pricing under Habit Formation and Catching Up with the Joneses," *American Economic Association, Papers and Proceedings*, May 1990, *80* (2), 38–42.

Ahmad, Yamin, "Money Market rates and implied CCAPM rates: some international evidence," *The Quarterly Review of Economics and Finance*, 2005, *45*, 699–729.

Atkeson, Andrew and Patrick J. Kehoe, "On the Need for a New Approach to Analyzing Monetary Policy," in "NBER Macroeconomics Annua," Vol. 23, NBER, 2009, pp. 389–425.

Backus, David K. and Gregor W. Smith, "Consumption and Real Exchange Rates in Dynamic Economies with Nontraded Goods," *Journal of International Economics*, 1993, *35*, 297–316.

Backus, David, Silverio Foresi, and Chris Telmer, "Affine term structure models and the forward premium anomaly," *Journal of Finance*, November 2001, *56* (1), 279–304.

Baier, Scott L. and Jeffrey H. Bergstrand, "Do free trade agreements actually increase members' international trade?," *Journal of International Economics*, 2007, *71*, 72–95.

Bansal, Ravi and Amir Yaron, "Risks for the Long Run: A Potential Resolution of Asset Pricing Puzzles," *Journal of Finance*, Aug 2004, *59* (4), 1481–1509.

_ and I. Shaliastovich, "Risk and return in bond, currency, and equity markets," Working paper, Duke University 2007.

_ , Dana Kiku, and Amir Yaron, "Risks for the Long Run: Estimation and Inference," 2007. Working Paper, University of Pennsylvania.

_ , _ , and _ , "An Empirical Evaluation of the Long-Run Risks Model for Asset Prices," *Critical Finance Review*, 2012, *1*, 183–221.

Beeler, Jason and John Campbell, "The Long-Run Risks Model and Aggregate Asset Prices: An Empirical Assessment," Working Paper 14788, NBER 2009.

Brandt, Michael W., Cochrane John H., and Pedro Santa-Clara, "International risk sharing is better than you think, or exchange rates are too smooth," *Journal of Monetary Economics*, 2006, *53*, 671–698.

Campbell, John Y. and John H. Cochrane, "By Force of Habit: A Consumption-Based Explanation of Aggregate Stock Market Behavior," *Journal of Political Economy*, April 1999, *107* (2), 205–251.

_ **and Robert J. Shiller**, "The Dividend-Price Ratio and Expectations of Future Dividends and Discount Factors," *Review of Financial Studies*, Winter 1988, *1* (3), 195–228.

Canzoneri, Matthew B., Cumby Robert E., and Behzad T. Diba, "Euler equations and money market rates:a challenge for monetary policy models," *Journal of Monetary Economics*, 2007, *54*, 1863–81.

Colacito, Riccardo and Mariano M. Croce, "Risks for the Long Run and the Real Exchange Rate," *Journal of Political Economy*, 2011, *119* (12), 153–181.

Epstein, Larry G. and Stanley Zin, "Substitution, Risk Aversion, and the Temporal Behavior of Consumption Growth and Asset Returns I: A Theoretical Framework," *Econometrica*, July 1989, *57* (4), 937–969.

Hodrick, Robert J. and Edward C. Prescott, "Postwar U.S. Business Cycles: An Empirical Investigation," *Journal of Money, Credit and Banking*, February 1997, *29* (1), 1–16.

Mehra, Rajnish and Edward C. Prescott, "The Equity Premium: A Puzzle," *Journal of Monetary Economics*, March 1985, *15* (2), 145–161.

Prais, G. J. and C.B. Winsten, "Trend estimators and serial correlation," Discussion Papers 383, Cowles Commission 1954.

Razin, Assaf and Efraim Sadka, *Foreign direct investment: analysis of aggregate flows*, Princeton University Press, 2007.

Reynard, Samuel and Andreas Schabert, "Modeling Monetary Policy," Working Papers 2010-04, Swiss National Bank 2010.

Rotemberg, Julio and Micheal Woodford, "Interest rate rules in an estimated sticky-price model," Working Paper series 6618, NBER 1998.

Tallarini, Thomas D., "Risk-sensitive real business cycles," *Journal of Monetary Economics*, June 2000, *45* (3), 507–532.

Weil, Philippe, "The Equity Premium Puzzle and the Risk Free Rate Puzzle," *Journal of Monetary Economics*, November 1989, *24* (3), 401–421.

Figure 1: Implied real risk free return vs *ex-ante* real federal funds rate - Power utility case

This figure plots the risk free return (r_t^*), resulting from the estimation of the eq.(4) with US data, against the *ex-ante* real federal funds rate (r_t), which has been computed as the difference between the nominal fed funds rate and the expected inflation from the VAR (18).

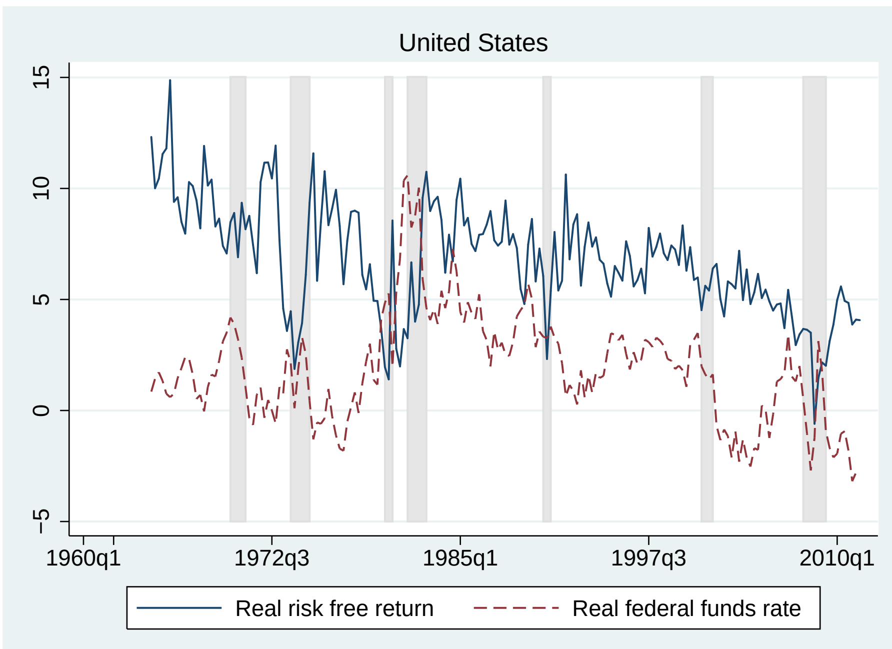

Figure 2: Implied nominal risk free return vs federal funds rate - Power utility case

This figure plots the nominal risk free return, estimated from equation (5) using US data, against the nominal federal funds rate.

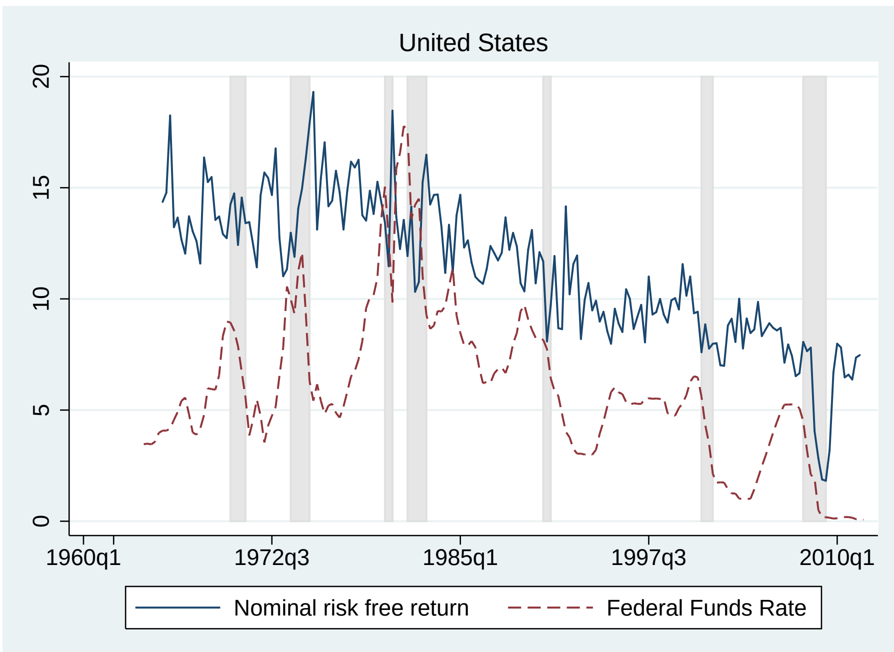

Figure 3: Implied real risk free return vs *ex-ante* real LIBOR - Power utility case

This figure plots the risk free return (r_t^*), resulting from the estimation of the eq.(4) with UK data, against the *ex-ante* real LIBOR (r_t), which has been computed as the difference between the LIBOR and the expected inflation from the VAR (18).

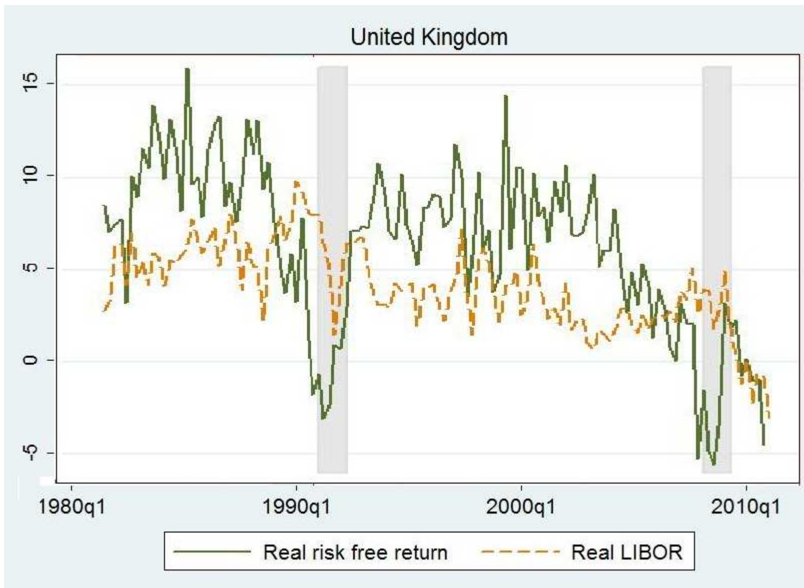

Figure 4: Implied nominal risk free return vs LIBOR - Power utility case

This figure plots the nominal risk free return i_t^* , estimated from equation (5) using UK data, against the LIBOR.

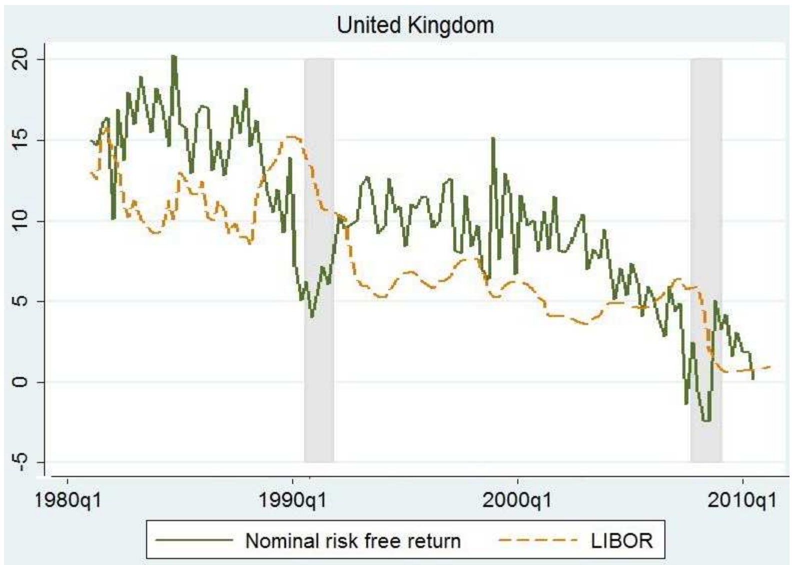

Figure 5: Implied real risk free return vs the *ex-ante* real federal funds rate - LRR model (Baseline)

This figure plots the real risk free return (r_t^*) implied by the baseline calibration of the LRR model (see Bansal and Yaron, 2004) for US data and compares it to the nominal federal funds rates.

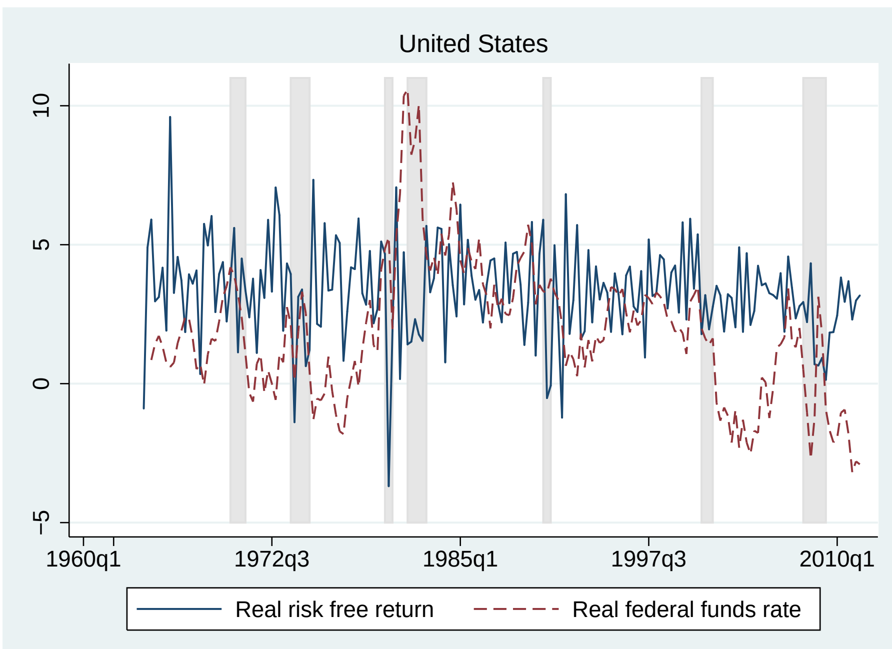

Figure 6: Implied real risk free return vs the *ex-ante* real LIBOR - LRR model (Baseline)

This figure plots the real risk free return (r_t^*) implied by the baseline calibration of the LRR model (see Bansal and Yaron, 2004) for UK data, and compares it to the *ex-ante* LIBOR.

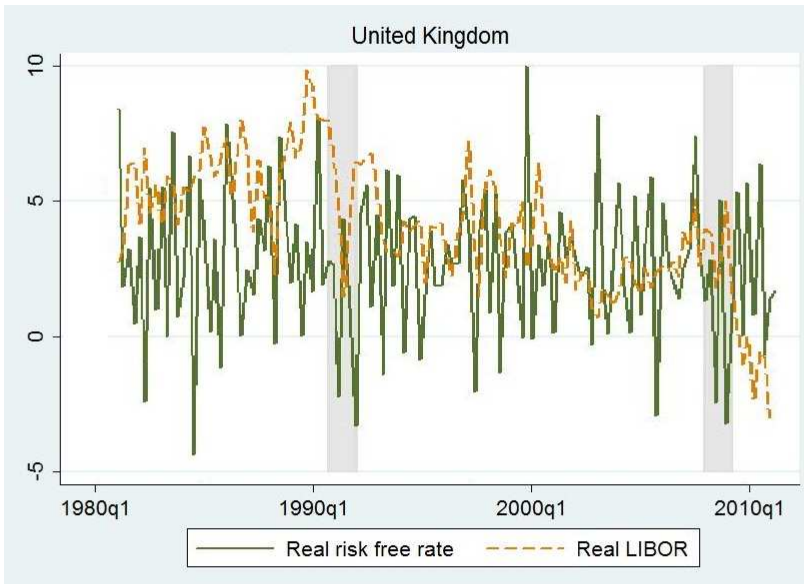

Figure 7: Implied real risk free return vs the *ex-ante* real federal funds rate - LRR model (Conditional volatility)

This figure plots the real risk free return (r_t^*) implied by the LRR model following the estimation procedure suggested by Bansal et al. (2007) and compares it to the *ex-ante* real federal funds rates.

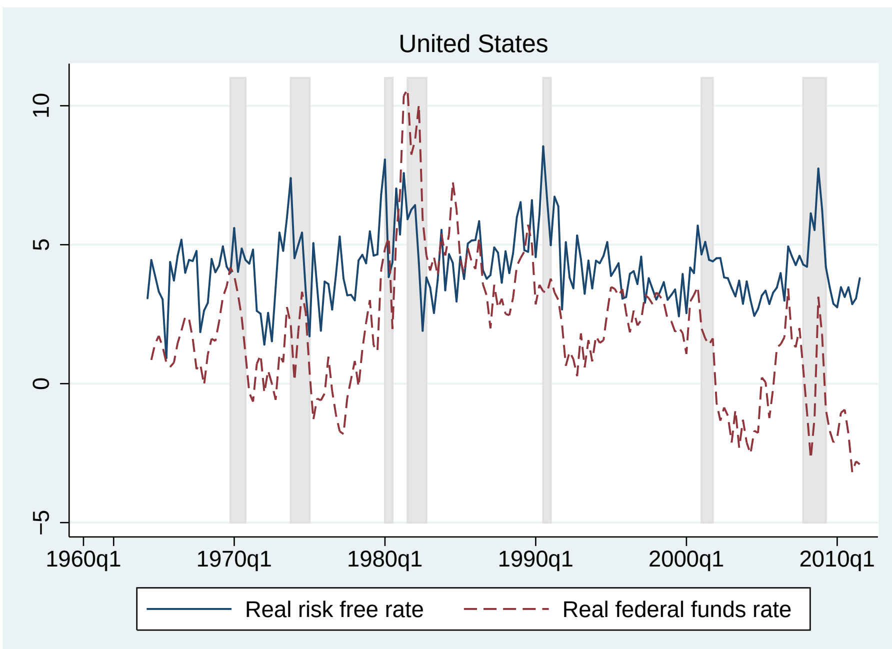

Figure 8: Implied real risk free return vs the *ex-ante* real LIBOR - LRR model (Conditional volatility)

This figure plots the real risk free return (r_t^*) implied by the LRR model following the estimation procedure suggested by Bansal et al. (2007) and compares it to the *ex-ante* real LIBOR.

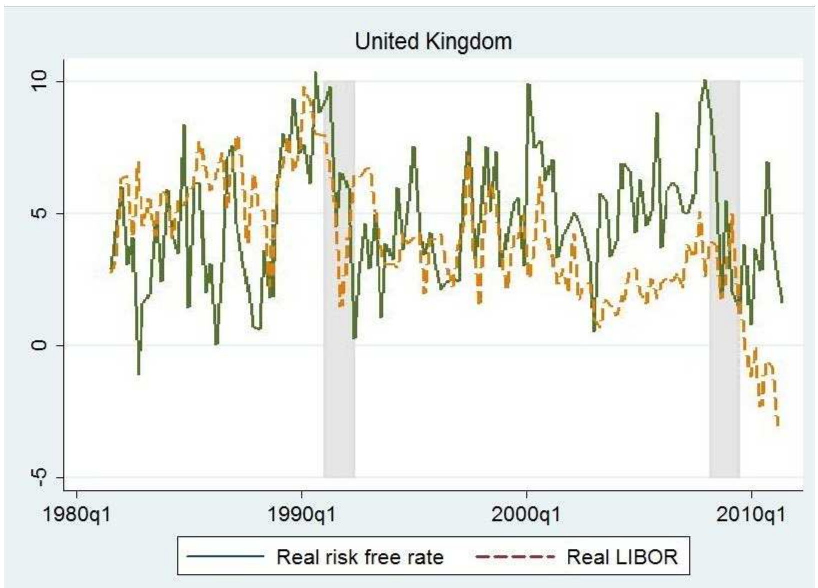

Figure 9: Exchange rate growth: model vs data - Power utility case

This figure plots the growth of the log-exchange rate resulting from (14), given the consumption dynamics implied by the standard power utility function. The model result is compared to the real dollar/sterling exchange rate growth obtained from the *Federal Reserve Bank of New York* (Website: http://www.newyorkfed.org/)

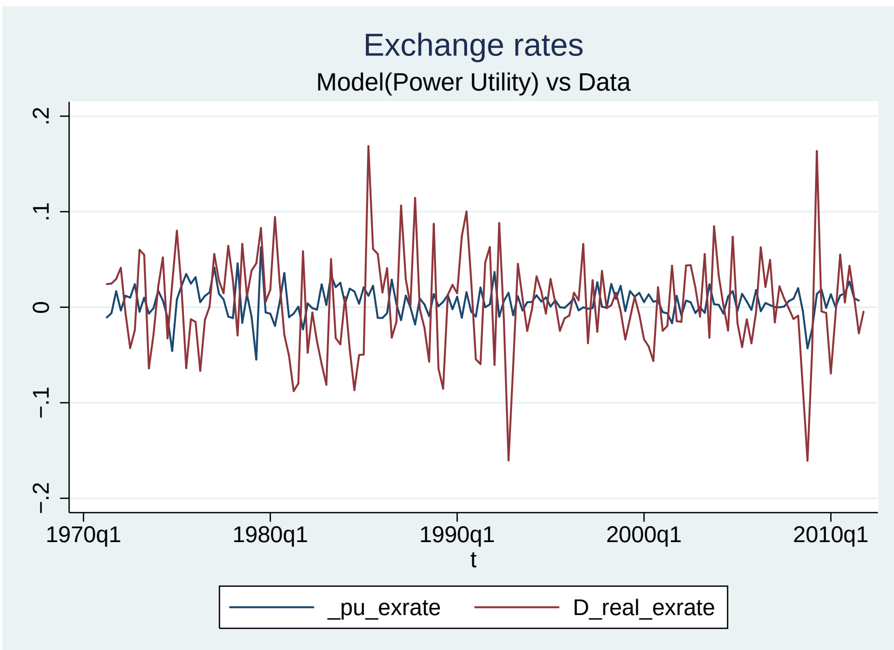

Figure 10: Exchange rate growth: model vs data - Restricted LRR model

This figure plots the growth of the log-exchange rate resulting from (14), given the consumption dynamics implied by the state space model in equation (24) and with the country-specific stochastic discount factors being estimated according to equation (17). The estimates are restricted to have complete symmetry between the two economies and fixing ρ_x and ρ_c to 1 and 0.3 respectively. The model result is compared to the real dollar/sterling exchange rate growth obtained from the *Federal Reserve Bank of New York* (Website: http://www.newyorkfed.org/)

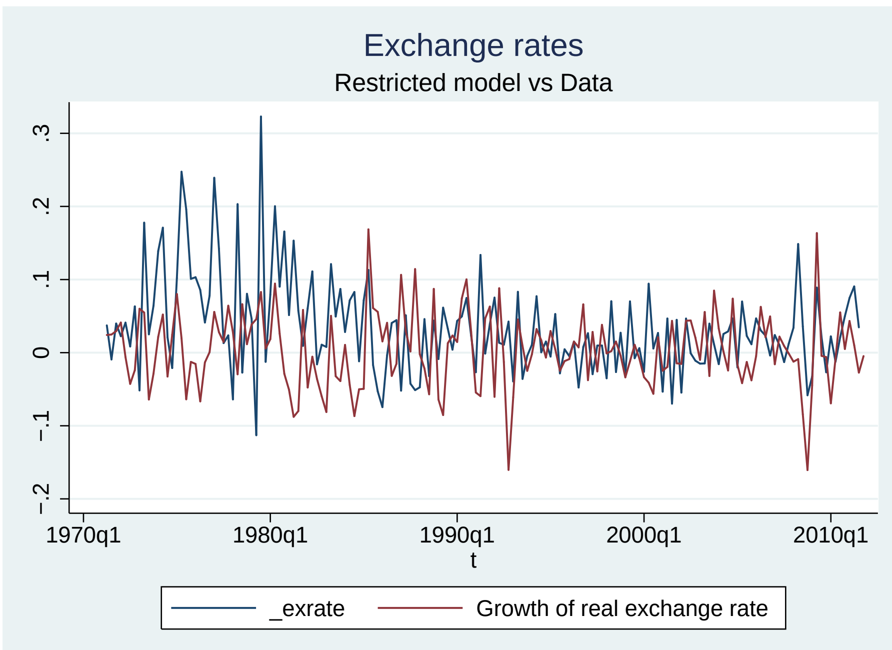

Figure 11: Exchange rate growth: model vs data - LRR model

This figure plots the growth of the log-exchange rate resulting from (14), given the consumption dynamics implied by the state space model in equation (24) and with the country-specific stochastic discount factors being estimated according to equation (17). The model result is compared to the real dollar/sterling exchange rate growth obtained from the *Federal Reserve Bank of New York* (Website: http://www.newyorkfed.org/)

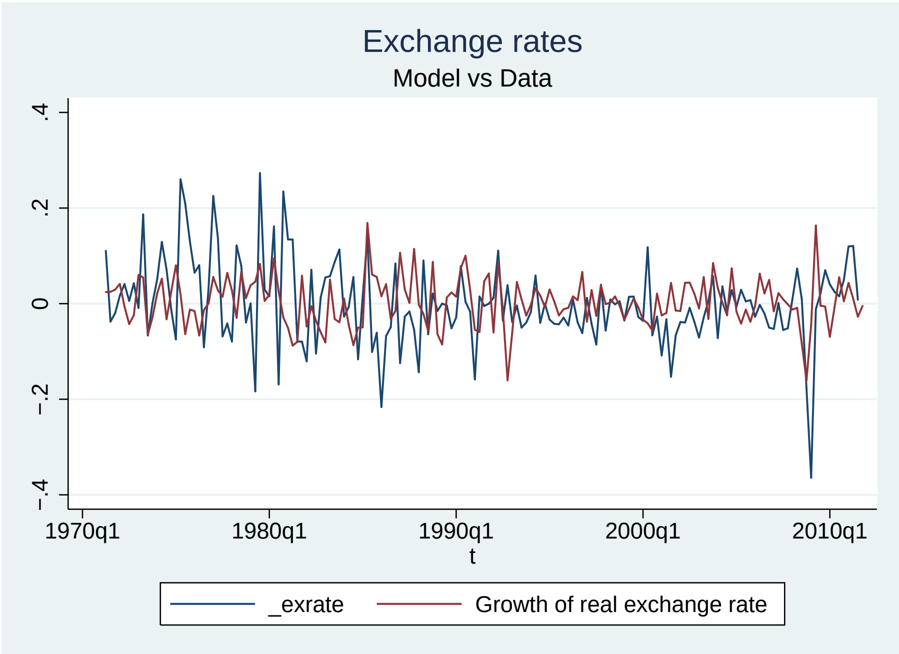

Table 1: Baseline calibration

The table shows the parameters used for the empirical analysis. The upper panel reports the parameters for the estimation of the risk free rate in equation (11). The model in (8)-(9)-(10) is calibrated at a monthly frequency. The lower panel collects the parameters for the estimation of the equation (17). The estimation of the model in (24) is based on quarterly data.

Panel A: Closed Economy estimates		
	United States	United Kingdom
Consumption dynamics		
ρ_x	0.919	0.848
σ	0.0023	0.0048
φ_e	0.205	0.126
μ	0.0016	0.0021
k_1	0.931	0.683
σ_w	6.52×10^{-7}	8.69×10^{-7}
v_1	0.988	0.898
Preferences		
β	0.998	0.998
γ	10	10
ψ	1.5	1.5

Panel B: Open Economy estimates			
	United States	United Kingdom	Restricted model
Consumption dynamics			
ρ_x	0.952	0.995	0.986
σ_c	0.0061	0.0142	1.68e-04
σ_x	0.0017	2.65e-04	4.77e-04
μ	0.013	0.0026	0.0086
k_c	0.965	0.319	0.642
Common parameters			
ρ_c^{hf}	0.143		0.3
ρ_x^{hf}	1		1
Preferences			
β	0.994	0.994	0.994
γ	4.25	4.25	4.25
ψ	2	2	2

Table 2: Summary statistics: model implied rates and money market rates - Power utility case

This table reports the means and the correlation coefficients between money market rates and the risk free returns implied by the model, estimated from equations (4) and (5).

	Real interest rates		*Nominal interest rates*	
	r_t	$r^*_{f,t}$	i_t	$i^*_{f,t}$
	(Data)	(PU)	(Data)	(PU)
United States				
Mean	1.88	4.70	5.89	9.070
Std. deviation	(2.476)	(2.455)	(3.498)	(3.215)
Min.	-3.184	-2.574	0.073	-3.135
Max.	10.590	12.591	17.78	16.957
Correlation		- 0.0219		0.4352***
United Kingdom				
Mean	4.08	2.37	7.47	9.73
Std. deviation	(2.365)	(4.720)	(3.881)	(4.896)
Min.	-3.016	-6.117	0.6	-2.858
Max.	9.840	16.155	15.79	20.110
Correlation		0.3342**		0.5667***

t-stat:* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 3: Interest rates spread and monetary policy - Power utility case

The table reports the main results from the following regressions:

$$(r_t^* - r_t) = \alpha + \beta_{r_t} r_t + \sum_{k=1}^{4} \gamma_k (r_{t-k}^* - r_{t-k}) + u_t$$

where the real spread between the real implied CCAPM rate r_t^* and the ex ante real federal funds rate r_t is regressed on its four lags and the federal funds rate r_t; and

$$(i_t^* - i_t) = \delta + \beta_{i_t} i_t + \sum_{k=1}^{4} \zeta_k (i_{t-k}^* - i_{t-k}) + e_t$$

where the spread between the nominal implied CCAPM rate i_t^* and the nominal federal funds rate i_t is considered instead. The relevant coefficients β_{r_t} and β_{i_t} for the federal funds rates (column 2) and for the LIBOR (column 3) are reported. To account for the autocorrelation in the error terms u_t and e_t, a Prais and Winsten (1954) procedure is used.

	United States	United Kingdom
Sample	1964q1-2011q4	1981q1-2011q4
Real interest rates		
Coeff. (β_{r_t})	-1.452***	-0.264
Std. error	0.113	0.136
Adj. R^2	0.55	0.58
Obs	185	116
Nominal interest rates		
Coeff. (β_{i_t})	-1.522***	0.006
Std. error	0.158	0.113
Adj. R^2	0.46	0.59
Obs	185	116
t-stat: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$		

Table 4: Summary statistics: the LRR model

This table compares the federal funds rate (r_t) and the risk free rate (r_t^*) implied by the LRR model of Bansal and Yaron (2004) and its subsequent modifications.
Baseline (second column) collects the results obtained when the equation (11) is estimated according to the baseline calibration in table 1 where the long run component x_t is calibrated.
Conditional (third column) and *No conditional* (fourth column) result from the estimation procedure introduced by Bansal et al. (2007) described in Section 3.2. In these cases, x_t is directly recovered from the data. The results in *No conditional* abstract from the conditional volatility of consumption.

	r_t		$r^*_{f,t}$	
	(Data)	(Baseline)	(Conditional)	(No conditional)
United States				
Mean	1.88	3.31	4.20	3.87
Std. deviation	(2.476)	(1.766)	(1.267)	(1.276)
Min.	-3.184	-3.69	0.917	0.629
Max.	10.590	9.60	8.542	8.207
Correlation		-0.0261	0,4295***	0,4406***
United Kingdom				
Mean	4.08	2.71	4.76	4.76
Std. deviation	(2.362)	(2.778)	(2.434)	(2.435)
Min.	-3.016	-4.366	-1.103	-1.094
Max.	9.840	10.007	10.403	10.422
Correlation		0.0102	0.2010**	0.1988**

t-stat: $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 5: Interest rate spread and monetary policy - LRR model

The table reports the main results from the following regressions:

$$(r_t^* - r_t) = \alpha + \beta_{r_t} r_t + \sum_{k=1}^{4} \gamma_k (r_{t-k}^* - r_{t-k}) + u_t$$

where the real spread between the real implied CCAPM rate r_t^* and the ex ante real federal funds rate r_t is regressed on its four lags and the federal funds rate r_t; and

$$(i_t^* - i_t) = \delta + \beta_{i_t} i_t + \sum_{k=1}^{4} \zeta_k (i_{t-k}^* - i_{t-k}) + e_t$$

where the spread between the nominal implied CCAPM rate i_t^* and the nominal federal funds rate i_t is considered instead. The relevant coefficients β_{r_t} and β_{i_t} for the federal funds rates (column 2) and for the LIBOR (column 3) are reported. To account for the autocorrelation in the error terms u_t and e_t, a Prais and Winsten (1954) procedure is used.

	(Baseline)	(Conditional)	(No conditional)
United States			
Coeff. ($\beta_{\Delta r_{t+1}}$)	-0.849***	-0.661***	-0.657***
Std. error	0.072	0.067	0.067
Adj. R^2	0.83	0.69	0.69
Obs	185	185	185
United Kingdom			
Coeff. ($\beta_{\Delta r_{t+1}}$)	-0.934***	-0.752***	-0.753***
Std. error	0.084	0.128	0.128
Adj. R^2	0.67	0.34	0.35
Obs	118	118	118
t-stat:* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$			

Table 6: Summary statistics: Exchange rate growth

Panel A compares the observed growth of the real dollar/sterling exchange rate with the exchange rate growth implied by equation (14) for the power utility case (column 2) and for the LRR models (columns 3 and 4). *Panel B* compares the observed growth of the real dollar/sterling exchange rate with the exchange rate growth implied by equation (14) for the power utility case (column 2) and for the LRR models (columns 3 and 4) on the subsample 1990q1 to 2011q3.

	Data	Power Utility	LRR Model Restricted	LRR Model
Panel A 1971-2011				
Mean	0.003	0.005	0.035	0.000
Std. deviation	(0.051)	(0.015)	(0.066)	(0.087)
Min.	-0.161	-0.055	-0.113	-0.364
Max.	0.169	0.063	0.323	0.273
Correlation	-	-0.0615	0.0564	0.0884
Panel B 1990-2011				
Mean	0.002	0.005	0.018	-0.015
Std. deviation	(0.049)	(0.012)	(0.042)	(0.066)
Min.	-0.161	-0.055	-0.070	-0.364
Max.	0.164	0.063	0.149	0.121
Correlation	-	0.0892	0.131	0.2271^{*}

t-stat: $* \; p < 0.05$