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Home bias and the persistence of real exchange rates

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ABSTRACT

This paper investigates the half-life of real exchange rates after taking into account the impact of home bias. Empirical results indicate that the half-life of real exchange rates is in the range of 1.5 to 2 years for four out of five countries after controlling the impact of home bias. These results support Obstfeld and Rogoff's (2000, NBER Macroeconomics Annual) view that home bias is crucial in explaining the PPP puzzle.

1. Introduction

Since the seminal paper by Rogoff (1996), the purchasing power parity (PPP) puzzle has become an interesting research issue in empirical studies of international finance. Rogoff (1996) points out that in existing literature the speed of PPP reversion measured by a half-life is about 3 to 5 years, which is too long to be consistent with the prediction of sticky-price models, such as Dornbusch (1976).[1] Conventional wisdom based on a sticky price model is that PPP deviations are corrected by price adjustments that determine the PPP reversion rate. Given the high integration of economies across countries and the intensive application of information technology in business, Rogoff (1996) points out that it is difficult to understand why it takes 3 to 5 years for prices to adjust to their equilibrium. He then argues that the reasonable half-life of real exchange rates should be around 1 to 2 years. Therefore, the observed half-life is too long to be consistent with the implication from models with nominal rigidity being called the PPP puzzle.

A conventional half-life measure is based on the unit-root parameter or an impulse response function. These measures may not be appropriate since they neglect the uncertainty of point estimates (Rossi, 2003), the presence of bias associated with inappropriate aggregation across heterogeneous coefficients (Imbs et al., 2002), time aggregation of commodity prices (Taylor, 2001), downward bias in the estimation of dynamic lag coefficients (Choi et al., 2004), and non-linear dynamics of real exchange rates resulting from the existence of transaction costs or pricing to markets (Taylor, 2001, Lothian and Taylor, 1996).

Although the PPP puzzle can be explained if the above issues are taken into account, no article empirically examines the role of home bias in explaining the puzzle. Obstfeld and Rogoff (2000) indicate that the role of home bias, associated with international trade costs in goods markets, may explain a number of empirical puzzles in international macroeconomics.[2] Recently, Mylonidis and Sideris (2008) apply panel regression to test whether the home bias is a source of real exchange rate deviations away from its PPP equilibrium, for G-7 economies in the post-Bretton Woods era. Their results support long-run PPP and that the home bias effect decreases over time. However, Mylonidis and Sideris (2008) do not investigate whether controlling the impact of home bias on real exchange rates explains the PPP puzzle. Therefore, the purpose of the paper is to answer the above question.

This study estimates time-varying home bias based on a Kalman filter and then applies Jorda's (2005) local projections to construct an impulse response function (IRF) of real exchange rates based on lags

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[1] The half-life indicates how long it takes for the impact of a unit-shock on real exchange rates to dissipate by half, which is a summary measure of persistence.

[2] The issue of home bias in tradable goods can be traced back to the elasticity of substitution (Armington, 1969) between imported and domestic goods due to changes in the relative price of those goods. Blonigen and Wilson (1999) indicate that home bias plays a prominent role in explaining the degree of substitutability between domestic and foreign goods based on the study of Armington-type elasticity coefficients for 151 manufacturing industries.

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of home bias and real exchange rates. Conventional impulse response based on the moving average representation assumes that the true data generating process (DGP) is known, which may not be true since the true DGP is unknown to researchers. An advantage of Jorda's local projections is that it allows researchers to construct IRF without knowing the true DGP. In other words, the IRF from local projections is robust to model misspecification.

Based on data from five industrialized countries over the period of the recent float, this paper points out that the half-life of real exchange rates, after controlling home bias, is in the range of 1.5 to 2 years for most of the countries under investigation. The above result supports Obstfeld and Rogoff (2000) view that home-bias is crucial in explaining the PPP puzzle.

The rest of the paper is organized as follows. Section 2 briefly describes the empirical model and estimation methodology. Section 3 explains empirical results. Finally, Section 4 provides conclusions.

2. The empirical model

Mylonidis (2008) proposes a very simple two-country, new Keynesian model to examine the significance of home bias in explaining the deviation of PPP, in which individuals maximize consumption of tradable goods and the price index of traded goods is assumed to be Cobb–Douglas. Mylonidis (2008) derives the following equation:

$$pppdev_t = s_t + \gamma\left(p_{ft} - p_{ht}\right) \tag{1}$$

where pppdev represents deviations from absolute PPP, s is the nominal exchange rates (foreign currencies per dollar); p_f and p_h denote the logarithmic of foreign and domestic tradable goods' prices, respectively; and $\gamma = (u - u^*)$, in which u and u* denote the share of traded goods expenditure allocated to home and foreign produced traded goods, respectively. Assuming that PPP holds Eq. (1) becomes:

$$s_t = \gamma\left(p_{ht} - p_{ft}\right) \tag{2}$$

If γ is positive then home bias exists. Mylonidis (2008) assumes that home bias is constant over time which may be restrictive. This paper assumes that u and u^* are time-varying and hence Eq. (2) can be written as follows:

$$s_t = \gamma_t\left(p_{ht} - p_{ft}\right).$$

It is worth noting that γ_t is not observed empirically. This paper applies the Kalman filter to estimate the unobservable home bias, γ_t. The Kalman Filter, provided by Kalman (1960), is an efficient recursive estimator for the unobserved state of a linear dynamic system in the presence of measurement errors.[3] Let the structure of a linear model consisting of measurement and transition equations be given as follows:

$$s_t = \gamma_t\left(p_{ht} - p_{ft}\right) + u_t,\ u_t \sim N\left(0, \sigma_u^2\right), \tag{3}$$

$$\gamma_t = \rho_0 + \rho_1 \gamma_{t-1} + \varepsilon_t,\ \varepsilon_t \sim N(0, \sigma_\varepsilon^2) \tag{4}$$

where γ_t is the unobserved state variable at time t. Eqs. (3) and (4) are measurement and transition equations, respectively. After estimating γ_t with the Kalman Filter, this study then constructs the impulse response functions of real exchange rates, by controlling the influence of home bias, γ_t.

Impulse response functions are widely used in macroeconomics to assess the persistence of variables and are estimated from the Wold decomposition of a linear vector auto-regression (VAR) model (Lutkepohl, 1993). Conventional impulse response analysis based on the VAR is to invert the VAR model to a vector moving average (VMA) representation and then construct the impulse response function based on the VMA representation. However, the above procedure is justified only if the model coincides with the true data generating process. As pointed out by Jorda (2005), conventional impulse responses based on the (MA) representation have several problems. First, the lag length of a VAR model may be very large in order to produce reliable impulse responses. Second, the VMA representation of a VAR may not be unique and different impulse responses can be due to different invertibility assumptions. Because of the above restrictions on the VAR approach, Jorda (2005) introduces a new method to estimate impulse response functions that is robust to the mis-specification of the unknown DGP. Instead of extrapolating into increasingly distant horizons as is conventionally done with vector autoregressions, Jorda (2005) estimates local projections at each period of interest. Therefore, the impulse responses can be defined without reference to the unknown DGP, even when its Wold decomposition does not exist (see Gary Koop et al., 1996; Potter, 2000). Jorda's method is a natural alternative to estimate impulse responses from VAR.

According to Hamilton (1994), an impulse response can be constructed as follows;

$$IR(t, s, d_i) = E(\mathbf{y}_{t+s} | \mathbf{v}_t = d_i; \mathbf{X}_t) - E(\mathbf{y}_{t+s} | \mathbf{v}_t = 0; \mathbf{X}_t) \quad s = 0, 1, 2, \ldots \tag{5}$$

where the operator $E(\cdot|\cdot)$ denotes the best, mean squared error predictor; $\mathbf{y}_t$ is an 2×1 random vector; $\mathbf{X}_t \equiv (y_{t-1}, y_{t-2}, \ldots)$; 0 is the 2×1 vector of zero; $\mathbf{v}_t$ is the 2×1 vector of reduced-form disturbances; and d_i represents the structural shock to the ith element in y. Eq. (5) points out that the statistical objective in constructing impulse responses is to obtain the best, mean-squared, multi-step predictors. These predictors can be obtained by recursively estimating a model that appropriately characterizes the dependence structure of successive observations (Jorda, 2005). Consider the following projection equation that projects $\mathbf{y}_{t+s}$ onto the linear space generated by $(\mathbf{y}_{t-1}, \mathbf{y}_{t-2}, \ldots, \mathbf{y}_{t-p})'$:

$$\begin{aligned} \mathbf{y}_{t+s} = \boldsymbol{\alpha}^s + \mathbf{B}_1^{s+1}\mathbf{y}_{t-1} + \mathbf{B}_2^{s+1}\mathbf{y}_{t-2} + \ldots \\ + \mathbf{B}_p^{s+1}\mathbf{y}_{t-p} + \mathbf{u}_{t+s}^s; \quad s = 0, 1, 2, \ldots, h \end{aligned} \tag{6}$$

where $\boldsymbol{\alpha}^s$ is the 2×1 vector of constants, and the $\mathbf{B}_i^{s+1}$ are matrices of coefficients for each lag i and horizon s + 1. The local projections are the collection of h regressions in Eq. (6). The impulse responses from Eq. (6), based on the definition on Eq. (5), are:

$$\hat{\mathbf{IR}}(t, s, d_i) = \hat{\mathbf{B}}_1^s d_i, s = 0, 1, 2, \cdots, h \ \text{ and } \ \mathbf{B}_1^0 = \mathbf{I},$$

where I is an identity matrix of an appropriate dimension. In other words, $\hat{\mathbf{IR}}(t, s, d_i)$ can be constructed by recursively estimating Eq. (6).

Constructing impulse responses from Jorda's local projections is easy and these responses can be obtained by conducting univariate least squares regressions for each variable at different horizons. Jorda (2005) demonstrates that impulse response estimates from local projections are consistent. Moreover, he shows that the statistical inference of these response estimates can be performed using standard heteroscedastic and autocorrelation (HAC) robust standard errors, such as Newey–West standard errors. These HAC standard errors correct for the moving average terms that exist in forecast errors.

We realize the possibility of unit roots and co-integration for the macroeconomic variables in our data set. However, Lin and Tsay (1996) find, based on monthly financial and macro-economic data of six major economies, that direct forecasting outperforms VECM-based forecasting in the presence of unknown unit roots and co-integration – even though unit roots and co-integration are ignored in direct forecasting. This paper therefore applies Jorda's local projections data in levels.

[3] Kim et al. (2009) applies the method of Kalman filter to investigate the long-run purchasing power parity for southeast Asian countries.