Real Exchange Rates in the Long Run: Evidence from Historical Data[1]

V. Anton Muscatelli[2]
University of Glasgow

Franco Spinelli
Università di Brescia

Carmine Trecroci
Università di Brescia

11 July 2001

[1] The authors gratefully acknowledge the CNR for research funding. Anton Muscatelli also wishes to thank the Comitato Economisti per Brescia for funding the research visits during which this paper was written. We are grateful to seminar participants at the Banca d'Italia and CESifo Munich for their comments. The usual disclaimer applies.

[2] Corresponding author address: Department of Economics, University of Glasgow, Adam Smith Building, Glasgow G12 8RT, United Kingdom. Tel: +44 141 330 6363; Fax: +44 141 330 4940; email: V.A.Muscatelli@socsci.gla.ac.uk

Abstract

We present empirical evidence on the forces driving real exchange rates in the long run. Using data from three industrialised countries, we find support for the hypothesis that productivity and fiscal shocks matter. There is also evidence, however, that the impact of fiscal shocks only matters in the short and medium-run. In some cases fiscal shocks cause depreciations, and this is probably explained by the monetary accomodation of fiscal shocks. The traditional Harrod-Balassa-Samuelson effect of productivity on real exchange rates is also found to be reversed in some cases, which demonstrates the importance of the distributive sector in driving productivity gains.

JEL Codes: F31, E44, C32

1 Introduction

The determination of real exchange rates in the long run has been the subject of a vast theoretical and empirical literature. Recently this has been driven by research on Purchasing Power Parity (PPP), which shows that we have only a partial picture of why deviations from PPP are so persistent over time[1].

Clearly, monetary (nominal) shocks cannot provide an explanation for this persistence, unless one assumes an unrealistic degree of nominal rigidity (see Rogoff, 1995). Fiscal and productivity shocks are seen as the main driving forces behind the real exchange in the medium and long run. Harrod (1933) originally used international productivity differences to explain the pattern of deviations from PPP, and this explanation was formalised by Balassa (1964) and Samuelson (1964) into what has become known as the Harrod-Balassa-Samuelson hypothesis. The impact of fiscal shocks on real exchange rates has been analysed both in the context of Keynesian aggregative models (Dornbusch, 1976), and intertemporal models of the open economy (Obstfeld and Rogoff, 1997).

Our approach is to use a structural VAR (SVAR) modeling framework to evaluate the importance of different shocks on the real exchange rates over various time horizons. The SVAR approach to the open economy has been employed by Clarida and Gali (1994), and more recently by Prasad and Kumar (1997) and Rogers (1999). Our contribution extends the current literature in a number of ways.

First, we use long runs of yearly data (approximately 120 years) for three countries. This enables us to focus more effectively on medium and long-run effects compared to studies which focus on post-World War II data[2]. Second, it is well known that the nature and impact of real and nominal shocks on

[1] For recent surveys, see Froot and Rogoff (1995) and Rogoff (1995, 1996). A number of papers have shown that there does not seem to be a significant degree of mean reversion in real exchange rates (e.g. MacDonald, 1996). However, more recent evidence using long runs of data (e.g. for example Abuaf and Jorion, 1990, Diebold et al., 1991) or panel data sets (e.g. Frankel and Rose 1995, MacDonald 1995) show some significant degree of mean reversion, but with a slow degree of convergence to the mean. One of the problems with unit root tests is highlighted by Engel (2000), who shows that these tests can suffer from significance size biases. In a recent paper, MacDonald and Ricci (2001) provide evidence for a panel of 10 OECD countries that the distribution sector has an impact on the real exchange rate.

[2] Rogers (1999) uses a VAR to model the real dollar-pound exchange rate from the late 1880s. Our data set contains a larger set of countries and a wider range of results. For a historical analysis of the impact of different fiscal instruments during the classical gold standard, see Kaminsky and Klein (1994).

the real exchange rate has varied over time[3]. This is due both to shifts in the underlying production structure of the industrial economies (for instance the different relative growth of the distributive sector and other non-traded services), and to the different regimes which have characterised the international financial system. To allow for these structural and regime shifts, we allow for time-varying parameters in our VAR models. Our results demonstrate the much greater persistence of both nominal and real shocks in the post-World War II era, not only compared to the interwar years, but also compared to the classical gold standard period. Third, in contrast to previous studies, we do not find uniform support for the Harrod-Balassa-Samuelson (HBS) effect. In some instances we find that productivity shocks can cause a real exchange rate depreciation, and we show that this is consistent with certain intertemporal models of consumption and production. Finally, instead of imposing certain restrictions on the long-run properties of our VAR model, we analyse the relative persistence of fiscal and productivity shocks and find support for the hypothesis, which derives from intertemporal models, that fiscal shocks, like all pure demand-side shocks, should only have temporary effects on the real exchange rate.

The rest of the paper is organised as follows. In Section 2, we outline some of the basic theoretical results in this area, which are now part of standard open economy macroeconomic theory. These motivate our VAR analysis. In Section 3, we set out our VAR modeling approach and we describe our data set. In Section 4 we present and discuss our results. Section 5 concludes.

2 Determinants of the Real Exchange Rate

2.1 Trends in the Real Exchange Rate: Theoretical Underpinnings

In our empirical models we distinguish between monetary, fiscal, and productivity shocks to the real exchange rate. The impact of monetary shocks on real exchange rates is relatively well understood in theory. Monetary shocks generate a temporary depreciation of the real exchange rate, but in the long-run money is neutral in all Keynesian aggregative and monetary-approach models. The persistence of the real depreciation depends on the degree of nominal rigidity in wage and price-setting. Some intertemporal models (see Obstfeld and Rogoff, 1997) display long-run non-neutrality, as

[3] For instance, Rogoff (1995, 1996) and Muscatelli and Spinelli (1999) show that the persistence of deviations from PPP seems to be more marked post-1946.

monetary shocks induce work-leisure substitution through a wealth effect. But, as noted by Obstfeld and Rogoff, this result is not robust to the specification of the consumers' optimisation problem and would disappear in an overlapping generations (OLG) model. *A priori* one would expect monetary shocks to have no long-run impact on the real exchange rate.

Turning to the effect of productivity growth, the standard HBS effect is best illustrated in a simple two-country, two-good model, where one good is tradeable (T), and is the numeraire, and the other is non-tradable (NT). For simplicity, suppose that the production technologies in the two countries are similar, but the countries have differing factor productivities:

$$
\begin{aligned}
Y^T &= A^T F^T(K^T, L^T), Y^N = A^N F^N(K^N, L^N) \\
Y^{T*} &= A^{T*} F^T(K^{T*}, L^{T*}), Y^{N*} = A^{N*} F^N(K^{N*}, L^{N*})
\end{aligned}
\tag{1}
$$

where the production technology is assumed to display constant returns to scale, and the two sectors are characterised by perfect competition. The two factors of production, capital (K) and labour (L) are mobile between sectors. Whilst labour is immobile internationally, capital is perfectly mobile, so that the real interest rate is equalized across the world. Define the home and foreign price indices as[4]:

$$
P = (1)^{\beta}(p)^{1-\beta}, P^* = (1)^{\beta}(p^*)^{1-\beta} \tag{2}
$$

where p and p^* are the prices of the non-tradable good at home and abroad. In a two-good world, the relative price level (real exchange rate) between the two countries depends solely on the relative price of non-tradables:

$$
P/P^* = (p/p^*)^{1-\beta} \tag{3}
$$

Shifts in factor productivity differentials will affect relative prices. To see this, we can use the firms' first-order conditions for profit maximisation in both sectors, using the production technologies in (1). Log-differentiating, we find that:

$$
\frac{\dot{P}}{P} - \frac{\dot{P}^*}{P^*} = (1-\beta)\left[\frac{\varphi_{LN}}{\varphi_{LT}}(g^T - g^{T*}) - (g^N - g^{N*})\right] \tag{4}
$$

[4] This price index assumes a utility function for consumers which depends on an index of total consumption which in turn is a Cobb-Douglas function of the consumption of tradables (share β) and non-tradables (share $1-\beta$). Using an alternative linear-homogenous function for the consumption index would complicate the formulation of the price index but not change the results.

where α^{LN} and α^{LT} are, respectively, the shares of labor income in the non-traded and non-traded sectors[5], and $g^x \equiv \dot{A}^x / A^x$. Equation (4) illustrates the BHS effect: a faster productivity growth in tradables relative to non-tradables will generate a real exchange rate appreciation. A corollary of this is that a country which experiences faster productivity growth should experience an appreciation of its real exchange rate relative to slower-growing countries, as productivity growth in tradables is likely to be faster than in non-tradables.

Although the HBS result is quite robust, alternative theoretical models can reverse the positive link between productivity growth and the real exchange rate. Increasing the number of factors of production or goods makes the basic trade model more complicated. In order to obtain some clear results, restrictions have to be imposed on the relative factor intensities of the different sectors and/or on the structure of demand if more than two goods are produced or there are more than two factors of production. A counterexample to the standard HBS result is provided by Devereux (1999)[6].

Devereux notes that, in contrast to Japan over the period 1950-1986, many of the East Asian economies since the 1970s have combined rapid economic growth with non-appreciating, or depreciating real exchange rates. Devereux develops a model in which three goods or services are produced. As above there are tradables, non-tradable goods (used only for domestic final consumption), but in addition there are non-traded distributional services. These services are used for the distribution and consumption of traded goods[7]. The production technologies of the tradable and non-tradables goods are as in (1), but distribution services are produced (using both capital and labour) by a continuum of monopolistically competitive firms. There are therefore increasing returns to specialisation in the distribution services sector.

The reason why this model can produce a depreciating real exchange rate in a fast-growing economy is that, as the economy grows the distribution networks in the country experience productivity deepening which lowers the (domestic) price of traded goods[8]. Again, taking the wholesale prices of

---

[5] i.e. $\alpha^{LT} = wL^T / Y^T$ and $\alpha^{LN} = wL^N / pY^N$, where w is the wage rate. Recall that the two countries have identical technologies apart from the productivity shift terms.

[6] The role of the distribution sector has generally been ignored in adjusting for the impact of productivity on teh real exchange rate. It only receives a brief mention in recent contributions (see Burstein et al, 2000, and Obstfeld and Rogoff, 2000). As far as we are aware the only direct test of the impact of the distribution sector can be found in MacDonald and Ricci (2001).

[7] The model can be generalised so that distributional services are used in the consumption and production of non-traded goods without changing the key results.

[8] Recall that whilst the wholesale price of traded goods is set in international markets

tradable goods as the numeraire, we can write the aggregate price index for the economy as follows, instead of (2):

$$P = (p^T)^{\beta}(p^N)^{1-\beta} \tag{5}$$

where p^T is the retail price of tradables and p^N as before is the price of non-tradables. The retail price of tradables is a price sub-index given by the (world) wholesale price of tradables, which is the numeraire, and the price of distribution services, p^D:

$$p^T = (1)^{\gamma}(p^D)^{1-\gamma} \tag{6}$$

Substituting (6) into (5) and then, as before, computing the impact of productivity growth in the various sectors on relative goods prices given the production structure, we can compute the impact on the real exchange rate. For simplicity, we can focus only on productivity growth in the home country, keeping productivity levels in the foreign country constant, which means that the evolution of real exchange rate is given by $\dot{P}/P$. Consider the special case where g^D, the productivity growth in distribution services is at the same rate as in tradable productivity[9]: $g^D = g^T$. In this case we have:

$$\left(\dot{P}/P\right) = \left[(1-(1/\rho))(1-\gamma)\beta + \epsilon_{LN}(1-\beta)\right] g^T - \left[\epsilon_{LT}(1-\beta)\right] g^N \tag{7}$$

where ρ is a measure of the elasticity of substitution between different varieties of distributional services[10], and $0 < \rho < 1$. Given that the first term in the square bracket in (7) is negative, it is apparent that if $g^N < g^T$ then HBS effect might be reversed and there might be a trend depreciation. The mechanism, as explained above, is the greater efficiency in tradable consumption promoted by productivity growth in the distribution network.

---

and obeys the law of one price, the domestic retail price of traded goods depends on the price of distribution services.

[9] This restriction is not necessary in order to reverse the HBS effect. It can be reversed even if $g^D < g^T$, but the conditions under which a trend depreciation will be observed are more complicated.

[10] The consumption of distributional services is given by

$$D = \left(\int_0^N x(j)^{\rho}\,dj\right)^{1/\rho}$$

where $x(j)$ is the output of each distributional services firm, and N is the number of firms. See Devereux (1999).

Finally, let's consider the impact of a fiscal shock on the real exchange rate. In standard Keynesian aggregative models under perfect capital mobility (e.g. models based on Fleming-Mundell and Dornbusch, 1976), a fiscal expansion will generate a permanent real appreciation and a deterioration in the trade balance. However, intertemporal models produce rather different results because they take full account of the impact of international borrowing and lending on wealth. Adding government spending (financed through lump-sum taxes) to the demand side in the above models will essentially distort the consumers decision by withdrawing resources from consumption. As shown in Obstfeld and Rogoff (1997), a government spending shock biased towards non-tradables (which is the usual case considered) will increase the relative price of non-tradables (a real exchange rate appreciation) in the short run[11]. The effect is short-lived, however, in the presence of capital mobility, as a process of investment in the non-traded sector will eventually reduce the marginal product of capital in the non-traded sector[12]. The short-run real appreciation resulting from a government spending shock is a robust result. A fiscal shock can generate a short-run depreciation only in models with a large degree of price-stickiness (e.g. Obstfeld and Rogoff, 1997) where output is essentially demand determined. Finally, it should be noted that intertemporal models consider the impact of fiscal shocks in a context of international capital mobility: during periods of imperfect capital mobility the impact of a fiscal shock might differ markedly from that described in Obstfeld and Rogoff (1997): the immediate impact of a fiscal expansion is likely to be a current account deficit which must either be financed or which will generate a nominal depreciation, and hence a short-run real exchange rate depreciation.

2.2 Implications for Empirical Modeling

The discussion in Section 2.1 can be summarised as follows. We would expect to find the following pattern of responses of the bilateral real exchange rate,

[11] The model considered assumes that investment and disinvestment in capital does not occur instantaneously because of costs of adjustment, hence allowing us to distinguish between the short-run effects of fiscal shocks on the real exchange rate and the long run, where the price of non-tradables is determined solely by supply factors (see Obstfeld and Rogoff, 1996, Appendix 4B, and Chinn, 1997).

[12] It is conceivable that if government spending is directed towards public investment (infrastructure), then it might increase the marginal product of capital in both the tradable and non-tradable sectors and increase output and consumption. In this case a fiscal expansion might have a permanent effect on the real exchange rate. This possibility is generally ignored in intertemporal open economy models. There also would seem to be little empirical support for a long-run effect, as we shall see below.

q, to a relative money supply shock $(m - m^*)$, a government spending shock $(g - g^*)$ and a productivity shock $(a - a^*)$:

Table 2.1: Impact on q

Shock:	$(m - m^*)$	$(g - g^*)$	$(a - a^*)$
Short run	+	+ (mobile/immobile capital)	?
Long run	0	0	- (HBS)/+ Devereux (1999)

The question mark on the short-run effect of a productivity shock is because the analysis in Section 2.1 highlighted the impact of productivity on the real exchange rate in the presence of factor mobility across sectors. These are long-run effects, and we would not anticipate a clear short-run impact of a productivity shock. Certainly there might be some short run response, but its sign is difficult to predict, a priori. Forward-looking expectations in foreign exchange markets might cause the nominal exchange rate to appreciate in response to an unanticipated productivity shock if economic agents expect an HBS effect to prevail. However, an increase in productivity might also trigger an increase in consumption as consumers anticipate the positive impact on wealth; this increase in consumption might cause a temporary current account deficit if it anticipates the investment (see Frenkel and Razin, 1995).

The other key points which emerge from our discussion in Section 2.1 are the following. First, the impact of both fiscal and productivity shocks might differ depending on the underlying productive structure and the prevailing regime in the international financial system[13]. For instance, we know that in a regime of immobile capital, the fiscal shock might cause a different short-run response in the real exchange rate compared to a regime of highly mobile capital. Also, we know that an industrial economy which is still developing its distributive networks might react differently to a productivity shock. More generally, as the impact of the fiscal and productivity shocks to the real exchange rate occurs through the supply-side of the economy, it is entirely possible that, as the productive structure changes (as is likely to be the case over our long sample period), the transmission mechanism will change over time. Both the sign and size of the fiscal and productivity multipliers might change through time.

A second and related point is that we might observe a persistent trend in the real exchange rate in some sample periods followed by other periods of

[13] As well as slowly-changing variables such as demographic factors and consumer preferences, which we do not consider here.

mean-reversion. This suggests a fundamental weakness in unit root tests on real exchange rates: the results are likely to be entirely sample-dependent. It also suggests a modeling strategy which (i) allows for some time-variation in the responses of the real exchange to the exogenous shocks in the system and (ii) avoids imposing long-run restrictions on the model, which incorrectly characterises its long-run properties.

3 Model Specification and Estimation

3.1 VAR Specification and Identification

We first turn to the issue of how to identify the structural disturbances in the VAR. Previous attempts to conduct a structural VAR analysis for the real exchange rate (e.g. Clarida and Gali, 1994, Lee and Chinn, 1998) have tended to impose some restrictions on the long-run properties of the structural VAR. Clarida and Gali (1994) and Lee and Chinn (1998) use Blanchard-Quah (1989) type restrictions on the model to identify the structural shocks. Rogers (1999) notes however, following Faust and Leeper (1997), that imposing Blanchard-Quah infinite-order restrictions on a finite-order VAR can cause problems. These problems are compounded since typically the number of variables included in a VAR are limited, so that the 'structural disturbances' are typically aggregations of different shocks, each of which may affect the macroeconomic variables included in the VAR in different directions. Rogers addresses the problem by estimating a reduced-form VAR on a set of differenced variables, and then imposing some structure on the long-run moving-average coefficient matrix.

Our approach here is different. We do not impose any long-run restrictions on the properties of the model along the lines of Blanchard and Quah (1989). In addition to the problems identified by Faust and Leeper (1997), we wish to avoid making assumptions about the underlying economic structure. As set out in Table 2.1, although theory does suggest some (very few) theoretical restrictions (e.g. money neutrality), in a number of cases the long-run impact of the productivity and fiscal shocks on the real exchange rate is not known a priori. We also wish to avoid any imposition of restrictions to examine whether the dynamic properties of our model are robust to different assumptions regarding the long-run relationships between the variables. For the same reason we avoid testing for (and imposing) cointegrating relationships between the real exchange rate and the explanatory variables: there is no reason to believe that such a cointegrating vector would be constant over the sample period.

Instead we focus on short-run restrictions. We perform a standard Cholesky factorisation to just-identify the structural shocks and calculate the impulse response functions. That is, we move from the reduced form of the VAR:

$$Z_t = c + B_1 Z_{t-1} + \dots + B_n Z_{t-n} + \epsilon_t \qquad E(\epsilon_t \epsilon_t') = \Sigma \tag{8}$$

to its structural form:

$$A_0 Z_t = c' + A_1 Z_{t-1} + \dots + A_n Z_{t-n} + \varepsilon_t \qquad E(\varepsilon_t \varepsilon_t') = V \tag{9}$$

where $B_i = A_0^{-1} A_i \ \forall i = 1:n$ and $\Sigma = A_0^{-1} V (A_0')^{-1}$; Z_t is the vector of k variables included in the VAR, ϵ_t is the vector of VAR innovations, and ε_t are the structural disturbances of interest.

This is done by imposing a lower triangular structure on the A_0 matrix[14]:

$$A_0 = \begin{pmatrix} a_{11} & 0 & 0 & 0 \\ a_{21} & a_{22} & 0 & 0 \\ a_{31} & a_{32} & a_{33} & 0 \\ a_{41} & a_{42} & a_{43} & a_{44} \end{pmatrix}$$

The vector of variables included in our VAR, and the corresponding vector of structural disturbances, is the following:

$$Z_t = \begin{pmatrix} (\frac{y}{N}) - (\frac{y}{N})^* \\ (\frac{G-T}{Y}) - (\frac{G-T}{Y})^* \\ \Delta((\frac{M}{P}) - (\frac{M}{P})^*) \\ q_t \end{pmatrix} \quad \varepsilon_t = \begin{pmatrix} \varepsilon_t^y \\ \varepsilon_t^g \\ \varepsilon_t^m \\ \varepsilon_t^q \end{pmatrix}$$

The vector of variables includes, in the following order: the differential in labour productivity between the home and foreign country, measured as output (y) per capita (N); the differential in the primary budget deficit, expressed as a proportion of GDP; the differential between the growth in real money balances in the two countries; and finally the real bilateral exchange rate, computed using aggregate price indices. The measure of productivity is the only one available over the long sample periods employed in our estimation. The chosen measure for the fiscal variable is a compromise: Rogers (1999) chooses G/Y as a measure of the fiscal stance, and we followed this approach in earlier work (Muscatelli and Spinelli, 1999). However, ignoring the role of taxation may be problematic if taxation changes have had

[14] This is obtained by factorizing Σ such that $\Sigma = PP'$, where $P = A_0 (V)^{1/2}$ (the Cholesky decomposition).

an important influence on the real exchange rate independently of government spending shocks, or if Ricardian equivalence does not hold[15], so that the macroeconomic response to an increase in government spending differs depending on how it is financed. Entering $G=Y$ and $T=Y$ separately would have addressed the first point, but it would have increased the dimension of the VAR, which is undesirable given the number of observations at our disposal. The monetary shock is captured in first difference form, following the approach in Rogers (1999). Unlike Rogers (1999) however, we have chosen not to break down monetary shocks into monetary base and money multiplier shocks. Although it is potentially helpful to understand whether monetary shocks originate from policy decisions of from the behaviour of the financial sector, we generally found that decomposing the monetary shock increased the dimensionality of the VAR without notably affecting the share of the total variance of q explained by the monetary shock[16].

Figures 1A-C plots the raw data[17]. It is apparent from this case that the trend in the real exchange rate q over some sub-samples can only be explained by the productivity variable, and that fiscal impulses can only have played a minor part[18]. This conclusion would not have changed if $G=Y$ had been used as the fiscal measure instead of the budget deficit. However, it should also be obvious that a simple HBS hypothesis linking productivity differentials to the real exchange rate will not do. There are periods of time when the productivity differential does not fully capture movements in the real exchange rate, and a time-varying parameter model is more likely to capture this feature of the data.

The ordering chosen for the variables implies the following short-run re-

[15] Muscatelli and Spinelli (1999) find that, in the case of Italy, Ricardian equivalence does not hold as the reliance on debt finance has had a significant impact on the real interest rate.

[16] It is well known that over the long historical period under scrutiny, the stock of real balances (even expressed as a difference between two countries) is not a stationary variable, and is unlikely to be an accurate measure of the relative stance of monetary policy (unlike G/Y). One alternative is to look at the velocity of money, but again that is not stationary over the sample period largely because there have been major changes in the financial structure, i.e. changes in the payments system and in the type of intermediation available.

[17] To take account of the war years, we pre-regress all the variables on a dummy variable for the First and Second World War. Especially in the case of Italy, accounting for the war years is imperative given the degree of economic dislocation which occurred.

[18] We do not present unit root tests for the individual series. As Engel (2000) shows, in the case of the real exchange rate there are serious size biases when unit root tests are employed. In addition, as discussed in Section 2, we would argue that the degree of integration of the series is unlikely to be invariant to the sample chosen because of regime shifts. Over different sub-samples our series might either be I(1) or I(0) depending on the policy regime.

strictions. The real exchange rate is assumed to react immediately to all the shocks, whilst the money variable reacts instantaneously to ε_t^g and ε_t^y, and the fiscal indicator only reacts to productivity disturbances. Productivity is only assumed to react with a lag to the other variables. The assumption that fiscal policy does not respond to exogenous shocks in monetary policy is defensible given that monetary policy generally accommodated fiscal policy over this historical period. Similarly, assuming that productivity does not immediately react to policy shocks is justifiable. The assumption that monetary policy does not react to exogenous real exchange rate shocks is slightly more problematic, especially during periods of fixed exchange rates, although it may be justified in part by noting that we are considering movements in the real exchange rate, rather than in the nominal exchange rate. Ultimately, as in the case of all structural VAR models, the validity of these just-identifying restrictions cannot be tested. However, we have tried alternative orderings for our variables (in particular reversing the order of the monetary policy variable and the real exchange rate, and reversing the order of the productivity and the fiscal policy variable[19]), and the conclusions remain basically similar[20].

3.2 Estimation Methods

As stressed in section 2.2, we would expect some variation over time in the response of the real exchange rate to the structural shocks. The impulse responses from the fiscal, monetary and productivity shocks are dependent on the production structure, consumer preferences and the degree of nominal rigidity and capital mobility. We therefore allow for some variation over time in the responses of the real exchange to the exogenous shocks in the system by estimating a time-varying parameter VAR. The estimation of VAR models with time-varying coefficients was pioneered by Doan, Litterman and Sims (1984). In this paper we follow this Bayesian approach to VAR estimation, which allows the parameters of the VAR to evolve as more observations are added[21]. The intuitive appeal of this approach is that policy regime and structural changes are modeled as a gradual evolution of the VAR coefficients (and hence the impulse responses).

[19] This is order to capture possible pro-cyclicality in the productivity variable given that it is defined as labour productivity.

[20] These alternative Cholesky decompositions are not included here for reasons of space, but are available from the authors on request. In future work, we plan to implement some of the robustness checks suggested in Faust (1998) to verify alternative identification schemes.

[21] The description here closely follows the notation used in Hamilton (1994), Ch. 13.

Consider the first equation of the reduced-form VAR:

$$z_{1t} = x_t' \beta_{1t} + \nu_{1t} \tag{10}$$

where $x_t = (1, Z'_{t-1}, Z'_{t-2}, \ldots, Z'_{t-n})'$ is the $(s \times 1)$ vector $(s = kn + 1)$ of regressors. The corresponding coefficient vector is:

$$\beta_{1t} = (c_1, \beta^1_{11,t}, \beta^1_{12,t}, \ldots, \beta^1_{1k,t}, \beta^2_{11,t}, \beta^2_{12,t}, \ldots, \beta^2_{1k,t}, \ldots, \beta^n_{11,t}, \beta^n_{12,t}, \ldots, \beta^n_{1k,t})'.$$

If we allow the VAR coefficients to vary over time, we need to specify the nature of the time-dependency. A state-space representation would have (10) as the measurement equation. We follow Doan et al. (1984) in postulating a Bayesian prior distribution for the first-period value of the coefficient vector $\beta_{11} \sim N(\bar{\beta}, P_{1|0})$. In addition, it is assumed that the VAR coefficients follow an AR(1) process; the transition equation of the system is therefore:

$$\beta_{1t} = (1 - \psi_1)\bar{\beta} + \psi_1 \beta_{1t-1} + \xi_{1t} \tag{11}$$

In (11), the parameter vector follows an autoregressive process, in which the weighting parameter ψ_1 determines the importance of the steady-state value for the coefficient vector. The disturbance term is uncorrelated with the disturbances in the original VAR: $cov(\xi_{1t}\nu_{1t}) = 0$, whereas the expected value $\bar{\beta}$ consists of a vector of zeroes with one as elements corresponding to the own variable at lag 1 $(z_{1,t-1})$ for each equation. This prior distribution holds that changes in the endogenous variable modelled are so difficult to forecast that the coefficient on its lagged value is likely to be near unity, while all other coefficients are assumed to be near zero. The prior distribution is independent across coefficients, so that the MSE of the state vector is a diagonal matrix.

The matrix $P_{1|0}$ is given by:

$$P_{1|0} = \begin{pmatrix} \vartheta \hat{c}_1^2 & 0' \\ 0 & (G \otimes C) \end{pmatrix} \tag{12}$$

where

$$G = \begin{pmatrix} \alpha^2 & 0 & 0 & \cdots & 0 \\ 0 & \alpha^2/2 & 0 & \cdots & 0 \\ 0 & 0 & \alpha^2/3 & \cdots & 0 \\ \vdots & \vdots & \vdots & \cdots & \vdots \\ 0 & 0 & 0 & \cdots & \alpha^2/n \end{pmatrix} \tag{13}$$

$$
C = \begin{pmatrix}
1 & 0 & 0 & \cdots & 0 \\
0 & w^2\hat{\sigma}_1^2/\hat{\sigma}_2^2 & 0 & \cdots & 0 \\
0 & 0 & w^2\hat{\sigma}_1^2/\hat{\sigma}_3^2 & \cdots & 0 \\
\vdots & \vdots & \vdots & \cdots & \vdots \\
0 & 0 & 0 & \cdots & w^2\hat{\sigma}_1^2/\hat{\sigma}_k^2
\end{pmatrix} \qquad (14)
$$

Also Q the variance-covariance matrix of ξ_{1t}, is given by: $Q = \pi_2 P_{1|0}$.

Doan et al. (1984) suggest the use of predefined values for the parameters in (12), (13) and (14). The following assumptions are made: $\alpha^2 = 0.07; w^2 = 1/74; \theta = 630; \pi_1 = 0.999; \pi_2 = 10^{-7}$; in addition $\hat{\sigma}_i^2$ is the estimated variance of the residuals for a univariate AR(n) regression estimated for series i. Note that the assumption is that the coefficient vector β converges only very slowly towards the mean. The factor γ defines the analyst's confidence that the first-order autoregressive coefficients β^1_{ii} relating z_{it} to z_{it-1}; is near unity for all i; θ is set sufficiently large to ensure that the prior expectation that the constant term is zero is given little weight; w^2 is set low to ensure that lags of other variables z_{jt} ($j \neq i$) are less useful in forecasting z_{it} than own-lags. Doan et al. find that these values work well for typical time series[22].

This general time-varying formulation involves forecasting the optimal state vector in each period, based on information available up to the previous period. Under the normality and independence assumptions about the disturbances, the computation of the state vector is simply obtained by applying the Kalman filter (see Hamilton, 1994). This allows us to obtain filtered estimates of the VAR parameters β and the residual variance-covariance matrix, V, for each observation in the sample. Orthogonalised impulse responses are finally computed according to the standard Cholesky decomposition, producing a set of different impulse responses for each period of our sample. Thus, for any time period in our sample, we will have a different set of responses of the real exchange rate to each of the shocks (fiscal, productivity and monetary) included in our model. The response to each shock will be based on estimates of the VAR which take account of the potential variation over time of the underlying structural parameters in the economy. This allows us to capture major regime shifts.

3.3 Data Employed

We employ the following data series. The price series used to construct the real exchange rates are consumer price indices. The productivity variable is

[22] We have experimented with alternative specifications, especially with regard to π_1, but find that our estimates are robust to values of π_1 between 0.3 and 0.999.

calculated as GDP at constant prices per capita. The budget deficit as the primary deficit as a proportion of GDP. The money stock series employed is a broad money (M2 equivalent) series for each country. The sample period used is 1868-1998, except in the case of the UK, where the sample starts in 1872. The pre-1950 GDP, population and price data were obtained from Flora (1983, 1987), Mitchell (1992, 1993), whilst data on exchange rates, fiscal variables and the money supply are compiled by Fratianni and Spinelli (2000) using a variety of historical sources. The data for the post-1950 period was checked for consistency using standard sources (IFS, OECD).

4 Estimation Results

4.1 US-UK

We first examine the impulse response functions for the real exchange rate between the UK and the US. Figures 2A-C show the impulse response of the real exchange rate following a productivity shock, a fiscal shock and a relative monetary shock in the UK for a number of periods after the shock (t+1, t+2, t+4, t+7 and t+10). Recall that, given our estimation method, we have a set of impulse responses for each sample period. Thus, Figures 2-4 show, for each time period, a set of impulse responses at different horizons when the VAR is estimated using data until that period. Standard errors are not shown in these figures as this would make the figures unclear. In fact for each period one can compute standard errors, using Monte Carlo bootstrap iterations of the baseline model (see Lutkepohl, 1991)[23]. We shall comment on the significance of the impulse responses as appropriate in our discussion.

Turning first to monetary policy (Figure 2C), we see that a monetary shock has the predicted response (a temporary depreciation), followed by a weaker response after 2-3 years. However, the monetary shocks are essentially insignificant, except for the inter-war float, and the mid-1980s. This is not surprising, given that in the whole of the post-1945 period, until the breakdown of Bretton Woods in 1972, the Pound was devalued just twice against the US Dollar. Hence monetary shocks were contained. However, perhaps surprisingly, even after 1972, monetary shocks are still not very significant, except for the mid-1980s. It appears that after the breakdown of Bretton Woods a monetary expansion generated a slim and very temporary real depreciation, one that only lasted 1-2 years.

The fiscal shock (see Figure 2B), as predicted, on impact appreciates the

[23] Standard error bands for ± 2 standard errors of the impulse responses for any single time period are available from the authors on request.

real exchange rate, and the effect appears to be quite persistent post-1950. The standard errors for the impulse responses show the appreciation following a fiscal shock to be significant for a period of about 3-4 years. As predicted by standard intertemporal open economy models (Obstfeld and Rogoff, 1997), fiscal shocks do not have permanent effects on the real exchange rate. The other interesting point to note is that the impact of fiscal policy seemed to be much less persistent in the pre-1945 years. Although a fiscal shock still tended to appreciate the real exchange rate in the period 1880-1940, the impact was less significant and more short-lived (2 years at most). Interestingly, the dichotomy does not seem to be between fixed and flexible-exchange rate regimes or between periods of high and low capital mobility: even during the classical gold standard the impact of fiscal shocks was short-lived. The main explanation for this seems to be that, with the exception of the two World Wars, the current account in the industrial economies were mainly driven by private consumption and saving behaviour and not by fiscal shocks. Note that the two World Wars and the period of increased fiscal spending from the mid-1930s mark the only phases when fiscal policy shocks impact significantly and persistently on the real exchange rate. The impulse responses just before and after the breakdown in Bretton Woods are similar, suggesting, quite surprisingly that the response of the nominal exchange rate post-1972 and relative prices pre-1972 in driving the real exchange rate following a fiscal shock was very similar.

Productivity shocks also seem not to have a significant impact on the real exchange rate until the acceleration in economic growth post-1950 (see Figure 2A). After 1950, there are signs of a significant HBS effect: the impulse responses show a progressive real appreciation, and even after 10 years the impact on the real exchange rate is sustained. This suggests that the HBS effect is a post-1945 phenomenon between the US and UK. By constructing standard errors we find that the effect is significant for horizons beyond 4 years, for the post World War II sample. The final point to note is that there seems to be no evidence from US-UK data that the importance of the HBS effect has diminished over time.

4.2 US-Italy

In the case of the real exchange rate between the US and Italy, the impact of a monetary shock is again seen to be short-lived (see Figure 3C). The standard errors show that, in fact the impulse responses are not significantly different from zero for the post-Bretton Woods float. During the Bretton Woods period, the depreciations are marginally significant because of the impact of the realignments. The impulse responses until 1913 are essentially insignificant. This is not suprising. The monetary shocks during the gold standard were relatively contained, and even though Italy abandoned the gold standard repeatedly during the period 1861-1913, between the years 1896-1913 it pursued a policy of monetary stability. The explanation for the fact that, during the inter-war years a monetary expansion led to a temporary appreciation is simple: Following the stabilisation of 1921, there was a period of gradual monetary contraction and the Lira returned to the Gold Standard in 1927. During these inter-war years, any monetary expansion fed through to domestic prices but not on the nominal exchange rate, thereby leading to a deterioration in competitiveness. The introduction of administered prices post-1935 exacerbated the position. Only with the abandonment of administered prices in 1943-46 did hyperinflation in Italy restore the real exchange rate to its equilibrium level. The low significance of the post-1972 impulse responses is more difficult to explain, except that the instability of the money demand relationships during this period may account for this: monetary growth was perhaps not a very good measure of the monetary stance at least until the mid-1980s in the case of Italy.

Turning now to the fiscal shock, in the case of the US-Italy exchange rate (see Figure 3B), a fiscal expansion is seen to lead to a temporary real depreciation in the post-1945 era. How can this be explained? There are two possible explanations. First, the approach we follow does not enable us to pick up a 'pure fiscal shock'. In the Italian case there is a very clear and unambiguous correlation between fiscal and monetary expansions. As demonstrated by Fratianni and Spinelli (1997, 2000), there is considerable evidence in favour of the 'fiscal dominance' hypothesis, whereby the Bank of Italy financed fiscal deficits through the creation of monetary base. In addition, much of Italy's investment in infrastructure in the post-1945 era was publicly funded, and hence we might be finding it difficult to disentangle the fiscal shock from a positive productivity shock due to public investment. Indeed, given the limited extent to which monetary shocks were important before 1965, the mixing of fiscal and productivity shocks is the most likely explanation for these results. Hence it is difficult, in the sense of Faust and Leeper (1997) to disentangle a 'pure fiscal shock' in the Italian case. Also, the Bank of Italy only gradually increased its effective independence from the fiscal authorities from 1981 onwards, and arguably only became fully independent in the run-up to entry to EMU in the 1990s. Second, in Italy's case the major fiscal expansion occurs post-1965, and hence the impulse responses are not very significant until that date. The standard errors for our impulse responses show that the results are more significant for the Bretton Woods period than post-1980. If one re-estimates the Bayesian VAR only on post-1945 data[24], one sees a major break between the pre- and post-1972 years, with latterly a more marked tendency for a fiscal shock to generate a temporary appreciation, as in the case of the US-UK real exchange rate.

The response of the real exchange rate to a productivity shock in the Italy-US case is of particular interest (see Figure 3A). Note that there is no evidence in favour of the HBS hypothesis, and indeed that a productivity shock generates a temporary depreciation for a period of up to 4 years (of which the first 2-3 years are significant, once one computes the standard errors). This is probably best explained in terms of the Devereux (1999) model: during the post-war era, Italy modernised its distributive and transport sector at the same time as its economy opened up, leading to a lower price of tradables. In essence we seem to observe in the Italian case an effect similar to that of the East Asian economies[25] in the period 1970-1995: faster productivity growth leading to no long-run real appreciation, and even a temporary real depreciation.

4.3 UK-Italy

The case of the UK-Italy real exchange rate is very similar to the US-Italy case (see Figure 4A-C). The monetary shock displays very similar properties. The fiscal shock shows, during the pre-1913 years, some tendency to generate a temporary real appreciation. But during the post-1950 period we again observe a persistent depreciation, which is probably due to an inability, as above, to disentangle the fiscal shock from the productivity shock, as public investment in tradables drove down tradable prices in Italy. However, again it should be stressed that the impulse responses to fiscal shocks are not very significant, except perhaps for a period in the late 1960s. The productivity shock shows again a tendency to generate a temporary depreciation (significant for the first two years), again providing support for the Devereux (1999) hypothesis.

4.4 Variance Decompositions for the Real Exchange Rate

The previous literature on modeling the real exchange rate using SVAR has produced contrasting results on the relative importance of monetary shocks.

[24] This is not reported here for reasons of brevity, but is available from the authors on request.

[25] Hence Italy seems to differ markedly from the case of Japan, which is usually cited as a leading example of the HBS effect (see Obstfeld and Rogoff, 1997).

Clarida and Gali (1994) find that for the G7 monetary shocks account for only about 2-10% of the variance of real exchange rate changes over all horizons. Rogers' (1999) study using UK-US historical data finds that monetary shocks account for between 19 and 60% of the real exchange rate variance, whilst fiscal and productivity shocks account for between 4 and 26%. The range of results emerge from a number of different specifications for his VAR.

In the case of our Bayesian VAR one can compute forecast error variance variance decompositions (FEVD) for any sample period. To better interpret the results we report the forecast error variance decompositions for a small number of years: 1910, 1930, 1960 and 1995. These are representative years during the classical gold standard, the interwar years, the Bretton Woods era, and the post-Bretton Woods era. The variance decompositions are reported in Table 4.1-3. It should also be noted that these are point estimates of the FEVD, and that in practice one can construct interval estimates.

In the case of the US-UK bilateral exchange rate, Table 4.1 shows that the productivity shocks only account for 5-6% of the FEV in the post-World War II period. The same cannot be said for the relative fiscal shock and, to a lesser extent, for the monetary shock. Indeed, for the four selected years, our indicator of budget balance shocks accounts for as much as 12 to 30% of the exchange rate's MSE, somewhat in line with Rogers' (1999) findings. The importance of monetary shocks seems to fade away in periods of flexible exchange rates, which may seem strange, but it basically illustrates the decoupling of monetary growth and exchange rate movements post-1980. The most likely explanation, as stressed by Rogers (1999), is instability in money demand.

Our results for the US-Italy case (Table 4.2) portray a rather different picture: here, productivity shocks seem to largely contribute (15% in 1995, 45% in 1910) to the forecast error variance of the bilateral exchange rate, particularly so in the Gold Standard years. However, the importance of productivity shocks has declined since 1945, and this confirms the absence of an HBS effect in the case of the US-Italy real bilateral rate.

Finally, Table 4.3 shows the decomposition computed for the UK Sterling-Italian Lira exchange rate. As in the US-UK case, the fiscal shocks seem to contribute significantly (28%) to the FEV, but only in the Gold Standard years. Monetary shocks, as noted in the case of the impulse responses, are less important, while productivity shocks explain about 3-4% of the FEV in the latter part of the sample.

4.5 Discussion

Our empirical results may be summarised as follows. First, we find that the evidence in favour of the HBS hypothesis is limited to one of the bilateral exchange rates in our study (the US-UK) case, and to the post-Bretton Woods era. The failure of the HBS prior to 1945 may be explained in various ways. First, as we noted in section 2, the HBS hypothesis is based on the assumption of international capital mobility. Second, the most dramatic acceleration in productivity growth came in the post-1945 era. Although the USA enjoyed considerable productivity gains during the period 1880-1940, the HBS effect might have been offset either by fast productivity growth in the distribution sector, or by the presence of considerable tariff barriers, which were only slowly dismantled after 1945. De facto the international trading system until the post-World War II era was dominated by regional preferential tariff schemes, which would have led to considerable distortions in relative prices.

Second, we find that the HBS effect is not present in the bilateral exchange rates between Italy and the US and UK. The absence of an HBS effect is probably attributable to the massive improvement in the distributive network in Italy during the post-World War II era. In this sense, Italy's experience seems to be similar to that of East Asia in the post-1970 period.

Third, our evidence shows strong support for the description of fiscal policy in standard intertemporal open economy models. Fiscal expansions in the US-UK case tend to generate temporary real appreciations, but there is no long-run impact on the real exchange rate. In the case of the US-UK and Italy-UK, the temporary loss in competitiveness following a fiscal shock is also detected during the classical gold standard, but to a lesser extent than in the post-1950 period. During the fixed exchange rate regimes the mechanism operates through the impact of fiscal policy on domestic demand and relative prices. Interestingly, post-1972 there appears to have been little change in the dynamics of the impact of fiscal policy on real exchange rates compared to the Bretton Woods era[26].

Fourth, as predicted by standard theory, monetary shocks only have a very temporary impact on the real exchange rate, lasting 1-3 years. Perhaps surprisingly, a similar trend is observed during fixed-rate regimes. During the Bretton Woods years this is probably best explained by the occasional realignments in the nominal exchange rate. This datum is also confirmed by the forecast error variance decomposition analysis, which also seems to

[26] In the case of Italy, it is difficult to disentangle the fiscal shock from the monetary and productivity shock in the post-1950 era. In general the fiscal shocks do not exert a significant impact on the real exchange rate.

confirm a recurrent unimportance of such shocks compared to the others we identified in explaining the real exchange rates' MSE. In the Italian case, the interwar years show a tendency for monetary shocks to lead to a loss in competitiveness as the nominal exchange rate was maintained at an increasingly unsustainable parity with gold.

5 Conclusions

In recent years, the theoretical literature on open economy macroeconomics has focused on the importance of fiscal policy and productivity growth in determining trends in real exchange rates. In part this has been driven by the observation that deviations from purchasing power parity are extremely persistent (see Rogoff, 1995, 1996). However, with notable exceptions (Rogers, 1999), most empirical studies tend to focus on the post Bretton Woods era. This makes it difficult to detect the impact of productivity trends, as predicted by the Harrod-Balassa-Samuelson hypothesis. It also makes it difficult to verify if, as predicted by standard intertemporal open economy macroeconomics, fiscal shocks do not have a long-run impact on the real exchange rate.

By using data for over 120 years for three industrialised economies in a structural VAR model, we provide a new perspective on these issues. We have shown that there is only limited support for the HBS hypothesis. However, as explained by a number of authors, notably Burstein et al. (2000), and Devereux (1999), the HBS hypothesis is based on a very particular theoretical model and rapid productivity growth may not necessarily lead to a faster growing domestic price level and an appreciating real exchange rate. Our empirical evidence in the Italian case shows considerable support for this alternative perspective to HBS. Our empirical exercise provides a useful adjunct to panel tests of the impact of productivity growth on the distributive sector and the HBS hypothesis (MacDonald and Ricci, 2001).

The other innovative aspect of our study is the use of a time-varying VAR to account for policy regime changes, and for shifts in the underlying production structure in these economies over the long sample periods studied. Our models show very clearly that the importance of fiscal and productivity shocks has increased since 1945. The use of time-varying models also allows us to point out the similarities of the dynamics of fiscal shocks in fixed and flexible exchange rate regimes. In addition, it clearly shows the very different operation of macroeconomic policy in the post-1945 era, with fiscal and monetary policy explaining a much greater proportion of the variability of the real exchange rate during Bretton Woods than during the classical gold standard. This provides an explanation for the observation in Rogoff (1995, 1996) that deviations from PPP have become more persistent in the post-World War II era.

References

Abuaf, N. and P. Jorion (1990). 'Purchasing power parity in the long run.' Journal of Finance, vol.45, pp.157-74.

Balassa, B. (1964). 'The purchasing power doctrine: a re-appraisal.' Journal of Political Economy, vol.72, pp.584-96.

Blanchard, O.J. and Quah, D. (1989). 'The dynamic effects of aggregate demand and supply disturbances', American Economic Review, vol. 79, pp. 655-673.

Burstein, A.T., Neves, J.C. and S. Rebelo (2000). 'Distribution costs amd real exchange rate dynamics during exchange-rate based stabilisations.' mimeo. Northwestern University.

Chinn, M.D. (1997). 'Sectoral productivity, government spending and real exchange rates: empirical evidence for OECD countries.' ch.7 in R. MacDonald and J. Stein (eds.) Equilibrium Exchange Rates. Amsterdam: Kluwer.

Clarida, R. and J. Gali (1994). 'Sources of real exchange rate fluctuations: how important are nominal shocks?' Carnegie-Rochester Conference Series on Public Policy, vol.41, pp.1-56.

Devereux, M.B. (1999). 'Real Exchange Rate Trends and Growth: A Model of East Asia.' UBC Discussion Paper n.99-05.

Diebold, F.X., Husted, S. and M. Rush (1991). 'Real exchange rates under the Gold Standard.' Journal of Political Economy, vol.99, pp. 1252-71.

Doan, T., R. Litterman and C. Sims (1984). "Forecasting and conditional projections using realist priori distributions". Econometric Reviews, 3(1), 1-100.

Dornbusch, R. (1976). 'Expectations and exchange rate dynamics.' Journal of Political Economy, vol.84, pp.1161-76.

Engel, C. (2000). 'Long-run PPP may not hold after all.' Journal of International Economics, vol. 51, pp.243-74.

Faust, J. (1998). 'The robustness of identified VAR conclusions about money.' Board of Governors of the Federal Reserve System, Discussion Paper n. 610, Washington DC.

Faust, J. and Leeper, E. (1997). 'When do long-run identifying restrictions give reliable results?' Journal of Business and Economic Statistics, vol.15, pp.345-53.

Flora, A. (1983). State, Economy and Society in Western Europe 1815-1975, vol. I, Campus: Frankfurt.

Flora, A. (1987). State, Economy and Society in Western Europe 1815-1975, vol. II, Campus: Frankfurt.

Frankel, J. and A. Rose (1995). 'A panel project on purchasing power parity: mean reversion within and between countries.' Journal of International Economics, vol. 40, pp.209-24.

Fratianni, M. and Spinelli, F. (1997). A Monetary History of Italy. Cambridge: Cambridge University Press.

Fratianni, M. and Spinelli, F. (2000). Storia Monetaria d'Italia (2nd ed.), Milano: ETAS.

Frenkel, J. and A. Razin (1995). Fiscal Policies and the World Economy. Cambridge MA: MIT Press.

Froot, K.A. and K. Rogoff (1995). 'Perspectives on PPP and long-run exchange rates.' In R.W. Jones and P.B. Kenen (eds.), Handbook of International Economics, Vol. 3,, North Holland, pp. 679-747.

Hamilton, J. (1994). Time Series Analysis. Princeton University Press, Princeton.

Harrod, R.E. (1933). International Econoics. Cambridge: Cambridge University Press.

Kaminsky, G. and M. Klein (1994). 'The real exchange rate and fiscal policy during the gold standard period: evidence from the United States and Great Britain.' NBER Working Paper n.4809.

Lee, J. and Chinn, M.D. (1998). 'The current account and the real exchange rate: a structural VAR analysis of major currencies.' NBER Working Paper n.6495.

Lutkepohl, H. (1991). Introduction to multiple time series analysis. Springer Verlag, New York.

MacDonald, R. (1995). 'Long-run exchange rate modeling: a survey of the recent evidence.' IMF Staff Papers, vol. 42, pp. 437-89.

MacDonald, R. (1996). 'What determines real exchange rates? The long and short of it.' Journal of International Financial Markets, Institutions and Money, vol. 8, pp. 117-53.

MacDonald, R. and Ricci, L. (2001). 'PPP and the Balassa-Samuelson Effect: the Role of the Distribution Sector' paper presented at CESifo Conference on Macro, Money and International Finance, Munich, May.

Mitchell, D. (1992). International Historical Statistics: Europe 1750-1988. Macmillan: London.

Mitchell, D. (1993). International Historical Statistics: The Americas 1750-1988. Macmillan: London.

Muscatelli, A., and F. Spinelli (1999). "Purchasing Power Parity and Real Exchange Rates: Do Productivity Trends and Fiscal Policy Matter?". University of Glasgow, Discussion Paper in Economics # 9921.

Obstfeld, M. and K. Rogoff (1997). Foundations of International Macroeconomics. Cambridge MA: MIT Press.

Obstfeld, M. and Rogoff, K. (2000). 'New directions for stochastic open economy models.' Journal of International Economics, vol. 44, pp.1-19.

Prasad, E. and Kumar, M. (1997). 'International trade and the business cycle.' mimeo. IMF, August.

Rogers, J.H. (1999). "Monetary shocks and real exchange rates'. Journal of International Economics, 49, pp. 269-288.

Rogoff, K. (1995). 'What remains of purchasing power parity?' Starr Center for Applied Economics, NYU Research Report n. 95-07.

Rogoff, K. (1996). 'The purchasing power parity puzzle.' Journal of Economic Literature, vol.34, pp. 647-68.

Samuelson, P. (1964). 'Theoretical notes on trade problems.' Review of Economics and Statistics, vol.46, pp.145-64.

Tables

US-UK	horizon		$Y \to q$	$G \to q$	$M \to q$	$q \to q$
1910	$h =$	1	0.0410	0.1797	0.0620	0.7170
		2	0.0410	0.1797	0.0621	0.7169
		3	0.0411	0.1797	0.0622	0.7168
		4	0.0412	0.1797	0.0622	0.7167
		5	0.0412	0.1797	0.0622	0.7167
		10	0.0413	0.1797	0.0622	0.7167
1930	$h =$	1	0.0004	0.3093	0.0252	0.6649
		2	0.0007	0.3092	0.0252	0.6647
		3	0.0019	0.3088	0.0252	0.6639
		4	0.0038	0.3082	0.0252	0.6626
		5	0.0058	0.3076	0.0251	0.6612
		10	0.0130	0.3054	0.0249	0.6565
1960	$h =$	1	0.0036	0.1972	0.0508	0.7482
		2	0.0056	0.1968	0.0506	0.7468
		3	0.0091	0.1961	0.0503	0.7443
		4	0.0142	0.1952	0.0499	0.7405
		5	0.0203	0.1940	0.0496	0.7360
		10	0.0578	0.1866	0.0477	0.7078
1995	$h =$	1	0.0033	0.1242	0.0073	0.8650
		2	0.0051	0.1240	0.0072	0.8636
		3	0.0082	0.1236	0.0070	0.8610
		4	0.0127	0.1231	0.0069	0.8571
		5	0.0183	0.1224	0.0069	0.8522
		10	0.0551	0.1178	0.0066	0.8203

Table 4.1. Forecast error variance decomposition, US-UK, 1872-1998.

The column header $X \to q$ refers to the proportion of forecast error variance of the bilateral real exchange rate q, h periods ahead, accounted for by innovations in variable X, where Y = differential in the productivity levels; G = differential in the relative fiscal stance; M = differential in relative nominal money growth (see main text for details).

US-Italy	horizon		$Y \to q$	$G \to q$	$M \to q$	$q \to q$
1910	$h =$	1	0.4510	0.0012	0.0108	0.5368
		2	0.4518	0.0012	0.0108	0.5360
		3	0.4525	0.0012	0.0108	0.5354
		4	0.4531	0.0012	0.0108	0.5347
		5	0.4535	0.0012	0.0108	0.5344
		10	0.4541	0.0012	0.0108	0.5338
1930	$h =$	1	0.1876	0.0002	0.0000	0.8119
		2	0.1860	0.0002	0.0000	0.8136
		3	0.1847	0.0002	0.0000	0.8149
		4	0.1836	0.0003	0.0000	0.8159
		5	0.1827	0.0003	0.0001	0.8168
		10	0.1802	0.0003	0.0001	0.8193
1960	$h =$	1	0.2026	0.0110	0.0044	0.7818
		2	0.2002	0.0111	0.0045	0.7840
		3	0.1974	0.0112	0.0045	0.7866
		4	0.1949	0.0113	0.0046	0.7890
		5	0.1929	0.0114	0.0046	0.7909
		10	0.1904	0.0115	0.0047	0.7932
1995	$h =$	1	0.1520	0.0050	0.0013	0.8414
		2	0.1492	0.0051	0.0013	0.8442
		3	0.1458	0.0051	0.0014	0.8475
		4	0.1426	0.0052	0.0014	0.8506
		5	0.1401	0.0052	0.0014	0.8531
		10	0.1367	0.0053	0.0015	0.8564

Table 4.2. Forecast error variance decomposition, US-Italy, 1868-1998.

The column header $X \to q$ refers to the proportion of forecast error variance of the bilateral real exchange rate q, h periods ahead, accounted for by innovations in variable X, where Y = differential in the productivity levels; G = differential in the relative fiscal stance; M = differential in relative nominal money growth (see main text for details).

UK-Italy	horizon	$Y \to q$	$G \to q$	$M \to q$	$q \to q$
1910	$h =$ 1	0.0021	0.2872	0.0174	0.6933
	2	0.0021	0.2871	0.0174	0.6933
	3	0.0022	0.2871	0.0174	0.6933
	4	0.0022	0.2871	0.0174	0.6933
	5	0.0022	0.2871	0.0174	0.6933
	10	0.0022	0.2871	0.0174	0.6932
1930	$h =$ 1	0.0698	0.0008	0.0002	0.9292
	2	0.0703	0.0008	0.0002	0.9287
	3	0.0709	0.0008	0.0002	0.9281
	4	0.0716	0.0007	0.0002	0.9274
	5	0.0722	0.0007	0.0002	0.9268
	10	0.0742	0.0007	0.0002	0.9248
1960	$h =$ 1	0.0701	0.0085	0.0725	0.9141
	2	0.0690	0.0085	0.0725	0.9151
	3	0.0678	0.0085	0.0725	0.9163
	4	0.0665	0.0085	0.0725	0.9176
	5	0.0654	0.0086	0.0725	0.9186
	10	0.0624	0.0086	0.0724	0.9116
1995	$h =$ 1	0.0435	0.0079	0.0090	0.9395
	2	0.0419	0.0079	0.0090	0.9410
	3	0.0399	0.0080	0.0090	0.9429
	4	0.0378	0.0080	0.0090	0.9450
	5	0.0358	0.0080	0.0090	0.9470
	10	0.0291	0.0081	0.0091	0.9535

Table 4.3. Forecast error variance decomposition, UK-Italy, 1872-1998.

The column header $X \to q$ refers to the proportion of forecast error variance of the bilateral real exchange rate q, h periods ahead, accounted for by innovations in variable X, where Y = differential in the productivity levels; G = differential in the relative fiscal stance; M = differential in relative nominal money growth (see main text for details).