Identifying Disturbances to Purchasing Power Parity

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I evaluate whether the equilibrium price of a freely traded currency is determined by purchasing power parity, subject only to disturbances confined in their magnitude and importance for real exchange rate variation. I present evidence from bivariate structural moving average representations that nominal exchange rates and relative aggregate price levels exhibit equal long-run responses following any permanent relative price level disturbance. Real exchange rates therefore display the long-run neutrality following such disturbances predicted by a monetary theory of the exchange rate. However, while these permanent price shocks account for almost all of the variance of relative price levels they explain a negligible fraction of real and nominal exchange rate variance. In fact, exchange rate variance is almost entirely attributable to large, significant and permanent shocks to nominal rates which are never reflected in relative price levels. Sources of variance in real and nominal exchange rates and in indices of relative national purchasing power are therefore orthogonal and there are departures from purchasing power parity at all horizons. The results are difficult to reconcile with models of permanent parity disturbances that rely on real shocks that alter the relative prices of home and traded goods, shocks that are commmonly attributed with producing at least transitional aggregate price level dynamics. Here, permanent PPP departures are never associated with transitory aggregate price level dynamics.

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"The fact that ... Purchasing Power Parity possesses such a remarkable stability is a sufficient reason for regarding (it) as the fundamental factor determining the rate of exchange and for classifying all other factors that may influence the rate...as factors of secondary importance, most suitably grouped under the heading of "disturbances"." - Gustav Cassel, 1928

In this paper, I present empirical results that clarify the conditions under which PPP can be said to have determined real and nominal exchange rates in both the short-run and long-run during the post-Bretton Woods era. The empirical models that I use allow me to discriminate between alternative economic interpretations of PPP, and suggest how one might reconcile the divergent evaluations of the validity of PPP observed in the extant literature.

Specifically, I estimate structural impulse response functions and variance decompositions that can characterize the short-run and long-run determination of real and nominal exchange rates and of relative aggregate price levels. The impulse response analysis shows that any permanent shock to relative price levels - such as would be produced by a shift in relative national money stocks - has exactly the effect for real and nominal exchange rates that would be predicted by a pure monetary theory of the exchange rate. Nominal exchange rates move one-for-one with relative price levels following these shocks, at all horizons, so that real exchange rates exhibit both short-run and long-run neutrality. The data therefore support PPP as a theory of how nominal exchange rates respond to permanent price level shocks, and suggest almost immediate convergence to equilibrium following such disturbances. This result conflicts strongly with the implications of "sticky-price", overshooting explanations of PPP departures.

However, the variance decompositions reveal that PPP fails as a complete account of exchange rate determination in the sense that these neutral price level shocks actually account for very little of the total variance of either real or nominal exchange rates. In fact, almost all of the variance of nominal exchange rates is accounted for by shocks that are entirely independent of relative price level movements, and it is precisely these nominal exchange rate shocks that drive most of the variation in real exchange rates. In the sample of this paper then, PPP fails due to large, significant and permanent nominal exchange rate shocks which are unrelated to the fundamentals that determine aggregate price levels - not to any short or long-run monetary non- neutrality.

Since the breakdown of the Bretton Woods fixed exchange rate regime, a voluminous literature has evolved that attempts to evaluate empirically whether PPP determines the relative price of freely traded currencies. In brief, the findings have been mixed in the sense that PPP has been rejected almost universally as a proposition about instantaneous real and nominal exchange rate determination, while alternative statistical techniques employed to identify long-run trends in exchange rates have produced divergent conclusions concerning the power of PPP as a theory of equilibrium exchange rate determination.[1]

In addition, interpretation of these empirical findings is clouded by the absence of a unique economic interpretation of PPP itself. Most simply, PPP is a proposition that the common currency price of an aggregate commodity (GDP or consumer good) basket will be equalized across spatially separated economies. As documented by Dornbusch (1987), this has been variously stated as a theory of tradable goods price determination under spatial arbitrage, or an approximation to exchange rate determination in the absence of frequent "real" disturbances, and as a monetary neutrality proposition.

How do these results clarify and contribute to the existing literature? First, they allow me to discriminate between two alternative economic interpretations of PPP and evaluate them independently. Second, they suggest statistical rationales for the divergent results obtained elsewhere in the literature.

PPP is frequently thought of as a proposition that international monetary disturbances are neutral for relative aggregate price levels denominated in a common currency. Shocks to one country's money stock should reflect equally and systematically in each component of the associated national price index and do not alter the relative prices of individual - tradable and non-tradable - commodities. They should therefore be reflected one-for-one in that country's relative aggregate price level, and so in the nominal exchange rate of the domestic country's currency. Alternatively, PPP is considered to be an empirical hypothesis that international goods market arbitrage dominates the determination of aggregate price levels, rather than

These results clarify some of the extant empirical literature. Observations on real and nominal exchange rates from the post-Bretton Woods era suggest that no form of instantaneous PPP holds. The data are characterized by volatile currency prices which diverge widely and persistently from the parity values implied by aggregate relative price series and this

PPP is therefore viewed at most as a long-run, equilibrium - rather than an instantaneous - condition for exchange rate determination, and this fact precludes inspection of raw data or computation of unconditional correlations between nominal exchange rates and relative price levels

[1] Empirical evaluations of equilibrium PPP that employ univariate tests are too numerous to document fully. Roll (1979), Meese and Singleton (1982), Adler and Lehmann (1983), Mussa (1986) and Diebold (1988) are among those studies that reject long-run PPP applying univariate non-stationarity tests to real exchange rates. By contrast, Diebold, Husted and Rush (1991), Cheung (1993) and Cheung and Lai (1993) reject the martingale hypothesis - fail to reject long-run PPP - in long spans of data when real exchange rates are modeled as more general long-memory processes. Huizinga (1987), Kaminsky (1988), and Grilli and Kaminsky (1991) find evidence of mean reversion in real rates using univariate variance ratio tests, while Abuaf and Jorion (1990) and Baillie and Bollerslev (1989) also support long-run PPP when cross-sectional information is incorporated into variance ratio statistics.

as a testing methodology. It requires identification of the long-run trends in exchange rates and prices.

Most generally, and following in essence Cassel (1928), equilibrium parity requires that any significant permanent disturbance to relative aggregate price levels across two countries should be ultimately and equally reflected in the nominal exchange rate between the currencies of those countries. All other disturbances to relative price levels and the nominal rate should be purely transitory for both variables or should account for a negligible fraction of their variance. Consequently, under PPP the ratio of any bilateral nominal exchange rate to relative price levels - the real exchange rate between two currencies - exhibits only limited and transitory deviations from a time invariant mean; there are limited disturbances to international parity in the common currency value of an aggregate commodity basket.

Applications to real exchange rates of univariate statistical methods designed to uncover violations of parity in the data have failed, however, to produce concensus on the validity of PPP doctrine. Yet tests of long-run PPP predicated on the univariate trend properties of real exchange rates have returned mixed results.[2] These tests use the fact that if long-run equilibrium currency values are determined by relative indices of national purchasing power, then their low-frequency behaviour will reflect this. Bilateral nominal exchange rates and relative aggregate prices should share a common stochastic trend, or cointegrate in the sense of Granger (1983), and their ratio - the real exchange rate - should be covariance stationary exhibiting purely transitory deviations from mean. The finding of a non-stationary real exchange rate then implies the failure of equilibrium PPP. It is now well-established that univariate tests for non-stationarity have (arbitrarily) low power to detect the presence of long-run parity in the data when the variance of permanent departures from PPP is small.[3] In addition, since univariate representations of exchange rates generally preclude identification of the empirical model with the multivariate economies within which long-run PPP is defined, any violations of equilibrium PPP uncovered using such methods cannot be attributed to specific economic phenomena. Univariate representations do not readily admit structural interpretation of PPP departures and therefore impart limited information for international monetary theory and for PPP doctrine in particular.

In this paper, I identify economies in which equilibrium PPP holds with a bivariate statistical model that satisfies a number of long-run neutrality properties. To evaluate the validity of PPP, I determine whether semi- restricted representations of this model, estimated from sample observations on nominal exchange rates and relative price levels, are consistent with these neutrality

[2] Froot and Rogoff (1994) describe the results of such univariate test applications to real exchange rates.

[3] See Quah (1992) and Cochrane (1991), and further discussion in Section II below.

properties. In particular, I impose on a reduced form some subset of the maintained long-run neutrality assumptions, and inspect the resulting representations for satisfaction of the remaining, over-identifying neutrality restrictions.

In fact, for PPP to be a good approximation to equilibrium exchange rate determination requires that two long-run neutrality properties hold for exchange rates and price levels. First, permanent shocks to relative aggregate price levels should be reflected equally and permanently in the nominal exchange rate, or any deviations from this condition must account for a negligible fraction of real exchange rate variance. Second, all other (by definition, purely transitory) shocks to relative price levels should be purely transitory for both variables and so induce insignificant real exchange rate variation at distant horizons. Both classes of economic disturbance should therefore exhibit long-run neutrality for the real exchange rate, irrespective of any transitory dynamics that they induce. In particular, under equilibrium PPP there can be no permanent nominal exchange rate movements that are associated with purely transitory relative price level dynamics and which account for a significant fraction of (long-run) real exchange rate variance.

I identify a bivariate structural moving average representation (MAR) which expresses nominal exchange rates and relative price levels as the outcome of current and historical realizations of permanent and transitory disturbances to relative prices. This is parameterized under the null as a reduced form bivariate autoregression of error-correcting form. I compute estimates of the structural infinite horizon multipliers for permanent and transitory relative price level shocks and impulse response functions for these shocks, and use forecast error-variance decompositions to determine the relative size of the implied transitory and permanent components in real exchange rates - the relative importance of transitory and permanent disturbances to PPP. I examine these statistics for evidence that the two long-run neutrality restrictions discussed above are satisfied.

I apply the representation to monthly G-7 data for the sample period 1975:1- 1991:12, computing the domestic spot price of foreign exchange using the $US as the numeraire (foreign country) currency, and conduct the analysis for both consumer and wholesale definitions of aggregate relative price levels to assess sensitivity of the results to the use of alternative data series reflecting different concentrations of tradable goods. The results that I obtain are as follows.

For no bilateral exchange rate according to the long-run neutrality criteria is long-run PPP satisfied. While permanent shocks to relative price levels typically are reflected one-for-one in the nominal exchange rate at distant horizons, and so exhibit the predicted long-run neutrality for the real exchange rate, these disturbances account for a small and insignificant percentage of both nominal and real exchange rate variance. By contrast, disturbances identified to be purely transitory for relative price levels engender large, significant and permanent movements in nominal and real exchange rates that account for almost all of the variance of these variables. Consequently, disturbances that are only transitory for relative prices cause permanent violations of PPP.

However, these disturbances are not significantly reflected in relative price variation at any forecast horizon; almost all of the variance of relative prices is accounted for by its "own" permanent disturbance. (In fact, relative prices have an insignificant transitory component, approximating a pure random walk.) Thus, for eleven of the twelve bilateral parity relations that I study, sources of variance in relative prices and in real and nominal exchange rates are orthogonal at all forecast horizons and permanent PPP departures are never reflected in relative price level dynamics.

These results are corroborated by a variety of robustness tests.[4] In addition, they are invariant to the choice of aggregate price index and small sample inference is invariant to the chosen ordering of the fundamental disturbances in the MAR.

The decomposition results that I report mirror the empirical observation that real and nominal exchange rates are approximately equally volatile. Importantly, they suggest that while the permanent component of prices is fully reflected in nominal exchange rates, as would be predicted by a monetary theory of the exchange rate, the predominant source of (permanent) real and nominal exchange rate variation is orthogonal to or "excess" relative to sources of variance in relative price levels. It cannot, therefore, be easily accounted for by real shocks to the relative prices of home and traded goods to which permanent PPP departures are commonly ascribed and to which are commonly attributed associated transitional aggregate price level dynamics. Here, in sample, permanent PPP disturbances are never reflected in aggregate price level movements.

I develop these points in the remainder of the paper as follows. I describe in Section I the maintained economic model and briefly review the relevant literature. In Section II I present a critique of univariate tests of this model, and develop in Section III an alternative mapping from the economic model to the bivariate empirical representations adopted here. Section III contains also the statistical criteria by which the hypothesis of equilibrium PPP is assessed and in Section IV I describe the results of applying these criteria to post-Bretton Woods G-7 data. Section V concludes.

[4]To evaluate robustness, I identify three alternative structural decompositions which should produce identical data representations under the maintained hypothesis. Each of these decompositions in fact produces a quite different representation to that observed in the sequel, suggesting specification error and so rejection of the null, and required long-run neutrality restrictions in these alternative representations are also consistently violated.

I. A Brief Review of PPP

In its most restrictive forms, PPP relies on international goods market arbitrage strengthened by necessary conditions on the cross-country similarity of aggregate price index components. Specifically, strong PPP asserts that under conditions of free trade, spatial arbitrage will ensure instantaneous equalization in spatially separated economies of the common currency price of an identical tradable commodity basket. Then, provided that different countries produce sufficiently similar and time-invariant aggregate commodity baskets, the nominal exchange rate between any two currencies should equal the ratio of the domestic to foreign price index pertaining to that basket. This assertion can be represented in logarithms as

$$\begin{aligned} e_t &= p_t^T - p_t^{T*} && (1a) \\ &= p_t + \theta - (p_t^* + \theta^*) && (1b) \\ &\simeq p_t - p_t^* && (1c) \end{aligned}$$

where e_t is the log of the domestic currency price of foreign exchange, p_t and p_t^* are the logs of the domestic and foreign aggregate goods price indices respectively, and $p_t^T = \theta + p_t$ and $p_t^{T*} = \theta^* + p_t^*$ are the log indices of aggregate tradable goods prices. Strong PPP therefore requires that the relative prices of tradable goods are both time-invariant and equal across countries.

Weaker forms of the doctrine, perpetrated *inter alia* by Cassel (1918), recognize that exact and instantaneous parity is unlikely to hold in the presence of frictions arising from spatial separation in international commodity markets. Wedges between bilateral nominal exchange rates and national indices of purchasing power, which are invariant over short horizons, result from impediments to free trade such as transportation costs, international transactions costs, and trade policy restrictions. Allowing for locational characteristics of tradable and, especially, non-tradable commodities suggests an additional sense in which proportional, rather than unitary, relations may be expected to hold between nominal exchange rates and relative aggregate price levels. Such characteristics conceivably may average to zero in the construction of p_t and p_t^*, but typically would preclude the equality of θ and θ^*. Despite these observations, "relative PPP" implies that common currency prices remain highly arbitraged and perfectly contemporaneously correlated, or

$$\begin{aligned} e_t &= c + p_t - p_t^* \Rightarrow && (2a) \\ \Delta e_t &= \Delta p_t - \Delta p_t^* && (2b) \end{aligned}$$

where c is a constant reflecting the effect of time invariant inter-location trade frictions and spatial characteristics of alternative country commodity baskets.

Casual inspection of data from the floating exchange rate regime since 1974 indicates that neither (1) nor (2) have much validity as explanations of nominal exchange rate movements over short horizons. Figure 1 plots monthly time-series of the six G-7 bilateral $US currency prices against relative wholesale price indices (WPI's), which reflect a high concentration of tradable goods. In this data, as is also true when (broader) consumer price indices (CPI's) are employed to construct relative price levels, movements in nominal exchange rates appear to little reflect movements in relative price indices month by month.[5] Consequently, real exchange rates - which are just deviations from strong PPP and given in logarithms by

$$r_t = e_t - (p_t - p_t^*) \tag{3}$$

- diverge widely from their mean (zero) values. Figure 2 plots demeaned log real exchange rates for the G-7 countries, where prices are measured both by wholesale and consumer indices, and illustrates these fluctuations clearly.

This empirical failure of instantaneous versions of purchasing power parity, which has been observed in all post World War I data, is accounted for by many models that admit transitory deviations from PPP but deliver parity as a long- run or steady state equilibrium condition. In the monetary tradition - of Dornbusch (1976) and Mussa (1982), for example - transitory PPP deviations are attributed to divergent speeds of adjustment in goods and asset markets following country-specific monetary shocks. Wages and prices are determined in markets characterized by menu and other costs of price adjustment and so respond slowly to monetary disturbances in contrast to nominal exchange rates - asset prices determined in frictionless spot markets. These divergent dynamics can produce over-shooting by nominal exchange rates of their new long-run values and induce potentially persistent deviations from PPP.

Implicit in many such traditional aggregative models, and an implication of neoclassical two-country, two-good models following Lucas (1982), is that steady state or long-run exchange rates conform to a PPP arbitrage condition provided there are no shocks that alter quilibrium relative prices of alternative goods. Equilibrium price levels and so the nominal exchange rate are determined by fundamentals reflecting money market equilibrium conditions - relative money stocks

[5] The data used throughout this paper are monthly, deterministically seasonally adjusted, logarithmically transformed and are demeaned prior to estimation. WPI's are averaged data from the International Monetary Fund's International Financial Statistics database, available from 1974:1-1991:12 for all but the Italian series for which the sample throughout the analysis is 1974:1-1989:12. Specifically, the WPI of Canada, France, Germany, Italy, Japan, the UK and the US respectively are: the aggregate industry selling price index; the price index of imported materials; the wholesale price index for industrials; a general wholesale price index; a general wholesale price index; the industrial output price index; and the price index of industrial output. CPI's are averaged data from the Citibase database, available from 1974:1-1991:12 for all countries. Nominal exchange rates are averages of daily noon spot rates for the price of one $US in terms of each country's currency units from the Citibase database, available 1974:1-1991:12 for all series.

and outputs. Specifically, both strands of the literature produce the familiar expression for an equilibrium nominal exchange rate (in logarithms)

$$\bar{e} = \bar{c} + \bar{p} - \bar{p^*} \tag{4}$$

$$\bar{p} - \bar{p^*} = (\bar{m} - \bar{m^*}) - (\alpha\bar{y} - \alpha^*\bar{y^*}) \tag{5}$$

$$\bar{c} = \bar{c}(\bar{p^T}/\bar{p}, \bar{p^{T*}}/\bar{p^*}) \tag{6}$$

where " $^-$ " denotes the long-run or steady state equilibrium value of the variable, $\bar{m}$ and $\bar{m^*}$ denote equilibrium money stocks in the two countries, and $\bar{y}$ and $\bar{y^*}$ represent equilibrium aggregate (traded and non-traded) output levels. Evidently, purely monetary and aggregate output shocks that permanently shift relative price levels have equal long-run effects for the nominal exchange rate and so are neutral for the equilibrium real exchange rate.

However, $\bar{c}$ reflects the equilibrium value of any cross-country differences in aggregate price index composition between home and traded goods, in preferences over alternative traded goods, and in production technologies across traded good sectors. It is apparent from (6) that there are permanent parity disturbances if $\bar{c}$ is not constant; movements in $\bar{c}$ permanently alter the equilibrium nominal exchange rate but cause only transitory or very limited movements in relative aggregate price levels. Shocks to preferences over alternative commodity baskets (as in Lucas (1982)) and permanent country-specific technology shocks within traded goods sectors (a la Balassa (1964) and Samuelson (1964)) are leading examples of such disturbances.

Cassel (1928) argues that $\bar{c}$ shifts are "abnormal deviations" from PPP exchange values and (if not purely transitory) are "...confined within rather narrow limits...", accounting for a negligible portion of real exchange rate variation. International economists and policy makers have consistently promoted such a Casselian "abnormal deviation" view to support long-run PPP as a good approximation to equilibrium nominal exchange rate determination or a good benchmark by which to judge current exchange rate valuation.[6] Thus, while two types of disturbance to PPP are identified in the extant literature - permanent disturbances to aggregate relative prices that are equally reflected in nominal exchange rates, and permanent nominal exchange rate shocks which are purely transitory for relative aggregate prices - under PPP the former are taken to dominate real exchange rate movements. Consequently, the real exchange rate should revert over time to a constant mean value following either permanent or transitory shocks to relative aggregate prices.[7]

[6]Robinson (1935) espouses a Casselian view that PPP is the proximate determinant of exchange rates, Mundell (1968, 1971) provide leading examples of the monetary theory of exchange rates and balance of payments which attribute exchange rate determination exclusively to money market disturbances that are neutral for relative prices, and the discussion in Dornbusch (1987) and Williamson (1983) present views of PPP as the appropriate benchmark for evaluating observed exchange rates.

[7]There are exceptions. The liquidity effects models of Grilli and Roubini (1992) and Schlagenhauf and Wrase

Mere inspection of the data or computation of unconditional correlations between nominal exchange rates and aggregate price levels cannot determine whether the real rate tends to revert to mean following individual disturbances to fundamentals. This observation is discussed in the context of existing tests of equilibrium PPP in the next section.

II. Univariate Representations of PPP

Relative price levels and the nominal exchange rate typically are characterized as linear stochastic processes driven by underlying shocks in the fundamentals which determine macroeconomic prices and quantities. In particular, the hypothesis of equilibrium PPP has been taken to imply that the real exchange rate is a covariance stationary stochastic process with time invariant and finite first and second moments. This is an appealing representation since it implies that innovations to PPP are finite-valued and finitely lived, and that following any innovation the real exchange rate eventually returns to a time-invariant constant (zero) mean value.[8]

The condition that r_t be covariance stationary is trivial if both e_t and $(p_t - p_t^*)$ are stationary processes. However, if stochastic fundamentals - embodied in production technology, preferences and the policy rule for monetary aggregates, for example - follow non-stationary processes which are reflected in non-stationarity of e_t and $(p_t - p_t^*)$, then long-run PPP requires cointegration of the nominal exchange rate and relative prices. Granger (1983) and Engle and Granger (1987) show that even if two (or more) variables are individually non-stationary, there can be a unique linear combination of them in which the non-stationary components are common and which is itself stationary as a consequence. Such variables are said to share a common stochastic trend or to be cointegrated. Under equilibrium PPP, any significant permanent shock to the relative aggregate price levels of two countries must be ultimately and equally reflected in the associated bilateral nominal exchange rate, and all other disturbances must be purely transitory or insignificant for the variance of both variables. Consequently, (log) nominal exchange rates and relative price levels should cointegrate with unique cointegrating vector $[1,-1]'$.

There is now a substantial body of evidence indicating that many macroeconomic time-series can

(1992) admit persistent deviations from parity attributable to asset market effects of monetary policy disturbances. Rogoff (1992) presents a model in which agents' intertemporal smoothing in traded goods' consumption produces intratemporal smoothing of traded and non-traded goods' prices and so a random walk real exchange rate. Betts and Smith (1994) present an economy in which there are permanent departures from PPP due to spatial separation of agents in international and domestic markets which prevents the cross-location exchange of goods. Froot and Rogoff (1994) discuss other models that exhibit non-parity equilibrium exchange rates.

[8] In fact, long-run PPP strictly requires only that the *first* unconditional moment of the real exchange rate be time invariant. Covariance stationarity is a stronger condition which has been imposed by empiricists and which has much intuitive appeal as an equilibrium condition.

be characterized as unit root processes, which is taken to imply that the underlying fundamentals are covariance non-stationary. Relative price levels and nominal exchange rates then may inherit this property and the received view is that this is the case.[9] This observation has generated many applications of univariate tests for covariance stationarity of r_t which test the null hypothesis that $\alpha = 1$ in the first order autoregressive moving average (ARMA) representation

$$r_t = c + \alpha r_{t-1} + v_t; \quad v_t = a(L)\epsilon_t. \tag{7}$$

Here, c is a constant, v_t is an MA process with $a(L)$ a lag polynomial satisfying conditions for invertibility, and the long-run equilibrium real exchange rate is defined as the unconditional mean of this process,

$$\bar{r} = (c/(1-\alpha)) = \bar{c}. \tag{8}$$

Evidently, strong PPP is violated whenever $r_t \neq \bar{r}$, while equilibrium PPP is violated whenever $\alpha \geq 1$.[10] When $\alpha = 1$, every shock to v_t has a permanent effect on r_t. When $\alpha < 1$, each shock to r_t is corrected at the rate of $(1 - \alpha)$ per period and so eventually dies out.

For the purposes of studying the dynamic properties of real exchange rates, one can also invert the univariate ARMA process (9) to generate an infinite order MAR

$$r_t = \bar{c} + d(L)\epsilon_t \tag{9}$$

where d(L) is a lag polynomial function with coefficients depending on α and a(L). The coefficients in d(L) summarize completely the dynamic behaviour of r_t and the persistence of PPP deviations from $\bar{c}$ in response to the single, reduced form shock ϵ_t.

A. Discussion

In many applications of non-stationarity tests to exchange rate data, researchers have been unable to reject the null of $\alpha = 1$ for nominal rates, relative prices *or* the real exchange rate and have concluded that equilibrium PPP is violated. Yet there are three important criticisms of the univariate approach which suggest its invalidity as a method for evaluating long-run parity.

First, it is well-documented (with Monte Carlo evidence) by Schwert (1987), Blough (1988), Gregory (1991) and others, that univariate non-stationarity tests have low power to discriminate between the null and close alternative hypotheses in small samples.[11] As discussed in Section I, it is

[9] See Meese and Singleton (1982) and Schotman (1989) on the case for non-stationarity of nominal exchange rates and Nelson and Plosser (1982) for evidence supporting non-stationarity of aggregate price indices.

[10] In fact, equilibrium PPP is also violated if either α or c is time-varying. These sources of violation are not considered in the current paper.

[11] Lo and MacKinlay (1989) show that similar results hold for variance-ratio tests of non-stationarity which also have been applied to real exchange rates.

precisely such borderline cases with which researchers since Cassel have been concerned. However, if the hypothesis of equilibrium PPP is correct, tests for unit roots have low power to detect this in the data.

More specifically, Cochrane (1991) and Quah (1992) show formally that any non-stationary process has both a permanent and a transitory component, the former having arbitrarily small variance, so that unit root tests have *arbitrarily* low power against some stationary alternatives in small samples. Yet, Section I suggests that a meaningful test of the hypothesis of equilibrium PPP relies on being able to identify the relative *size* of the permanent component in real rates - of $\bar{c}$. If a real exchange rate, for which the unit root hypothesis cannot be rejected, has a permanent component with small variance, then deviations from PPP may be primarily transitory despite the statistical non-rejection. As shown in (10), univariate representations allow derivation only of one, reduced form error process which may be a linear combination of multiple underlying disturbances some with permanent and some with transitory effects. Such representations cannot identify and measure alternative sources of PPP disturbance in the absence of additional information that imposes some structurally interpretable decomposition of variance for v_t.[12]

Finally, and most generally, without the means to identify multiple, structurally interpretable sources of real exchange rate disturbances univariate methods have little potential to inform theory with explanations for the failure of long-run PPP.

III. Multivariate Representations of PPP

Univariate evaluations of long-run real exchange rate behaviour fail to exploit additional information that is available in multivariate systems. This information can be used to identify multiple sources of disturbance to real exchange rates, to measure their relative size in accounting for the variance of real rates, and to characterize exchange rate dynamics that inform theory on the role of different fundamental shocks.

In particular, multivariate systems admit identification of both transitory and permanent components in exchange rates, as discussed - in a more general context - by Quah (1992) and Blanchard and Quah (1989).[13] In addition, King and Watson (1992) suggest that "structural" or permanent-

[12]A related criticism of univariate methods is that they may lead to at best uninformative and at worst misleading representations of real exchange rate dynamics. Cochrane (1991) emphasizes that there exist unit root processes (with permanent components having small innovation variance) whose autocorrelation and likelihood functions are arbitrarily close to those of given stationary series. For such processes, the asymptotic distribution theory derived under the false alternative may be a better guide for statistical inference in small samples than the non-standard (correct) distribution theory applicable under the assumption of non-stationarity. This challenges the use of unit root test results as information for specifying univariate representations of exchange rate dynamics in borderline cases.

[13]Quah (1992) shows that multivariate VAR representations can identify both the (orthogonal) permanent and

transitory VAR decompositions also can be invoked to evaluate propositions about long-run neutrality in the presence of non-stationarity in the endogenous variables of an economic model. The Quah-King-Watson methodology can be readily extended to a test of equilibrium PPP (despite the presence of cointegration under the null, which is a case not considered by King and Watson) since this hypothesis translates directly into long-run neutrality propositions concerning the effects of permanent and transitory relative aggregate price level disturbances for nominal exchange rates.

The next section shows how bivariate structural AR's can be used to identify and measure disturbances to real exchange rates, and to evaluate equilibrium PPP in this context.

A. Representation and Identification

The class of models which predict equilibrium PPP have been characterized above. These maintain that, while there are multiple disturbances to nominal and real fundamentals that drive aggregate price levels in each of two economies, only permanent aggregate relative price shocks drive long-run nominal exchange rate movements. All transitory aggregate price disturbances have approximately zero long-run effects for both variables. This implies that bivariate representations of bilateral nominal exchange rates and relative prices, which allow identification of only two disturbances, are "complete" provided that all of the underlying fundamental shocks to relative price levels (or the nominal exchange rate) can be aggregated into two orthogonal, representative innovations. Specifically, when nominal rates and relative price levels are subject to permanent shocks, a bivariate system in one aggregate permanent relative price disturbance and one aggregate transitory relative price disturbance can potentially represent all such maintained models. Some formalization of this idea follows, and a description of bivariate AR's that serve as representations of any abstract economy characterized by equilibrium PPP. [14]

Assume that the vector of non-stationary economic variables, $X_t = [e_t, (p_t - p_t^*)]\prime$, is jointly, completely and fundamentally determined by a fixed set of underlying disturbances which can be aggregated into a single transitory relative price disturbance, η_t^T, and a single permanent relative price disturbance, η_t^P, as $\eta_t = [\eta_t^T, \eta_t^P]\prime$. The variables in X_t are cointegrated with unique

transitory components of non-stationary series. The appeal of Quah's multivariate method is that it admits direct estimation of the two components in a simple VAR system. Such structural VARs have been proposed and applied by Blanchard and Quah (1989), Quah (1992) and Cochrane (1992) as alternative methods for estimating the permanent and transitory components in time series which are characterized by stochastic non-stationarity.

[14] Blanchard and Quah (1989) and Quah (1992), show that representations in which all permanent and all transitory disturbances are aggregated into single stochastic variables in this manner, must satisfy some regularity conditions to be valid. Specifically, two conditions must hold. First, all of the underlying permanent shocks must elicit very similar dynamic responses of the two variables. Second, the transitory disturbances can have different dynamic effects for the variables but must leave (nearly) unchanged the dynamic relation between them. It is assumed - consistently with the maintained hypothesis - that these regularity conditions are met.

cointegrating vector [1,-1]$\prime$ under the maintained economic model.

Then from Wold's Theorem the (2×1) jointly covariance stationary stochastic vector process, ΔX_t, has infinite (causal) vector MAR

$$\Delta X_t = \Gamma(L)\eta_t \tag{10}$$

where $\Delta X_t = [\Delta e_t, \Delta(p_t - p_t^*)]'$, $\Gamma(L)$ is a matrix lag polynomial summarizing the dynamics of the system and satisfying conditions for stationarity and invertibility, and η_t is a (2×1) vector of the orthogonal permanent and transitory fundamental disturbances for X_t, with variance-covariance matrix Σ_η. In addition, from the Granger Representation Theorem, we know that $\Gamma(1)$ has rank N-p=2-1=1, where N is the number of variables in the vector X_t and p is the number of independent cointegrating vectors between the elements of X_t.

Knowledge of the parameters in $\Gamma(L)$ and the orthogonal innovations η_t would provide a complete description of the system's dynamics. However, the MAR represented by (10) typically cannot be estimated directly, but must be identified from a reduced form parameterization which is consistent with the maintained economic model. From Engle and Granger (1987) we know that if the 2×1 vector X_t is non-stationary, but the linear combination given by r_t in (3) is stationary, then the covariance stationary exchange rate/relative price system can be parameterized in one of two ways.

First, and my focus in this paper, one can estimate a vector error-correction model (VECM) given by

$$\Delta X_t = -\gamma r_{t-1} + B(L)\Delta X_{t-1} + \epsilon_t, \tag{11}$$

where r_t is the real exchange rate and the "error- correction term" which captures the (maintained) common permanent component in e_t and $(p_t - p_t^*)$. Its inclusion restores the spectral density of the model to full rank, accounting for the singularity in $\Gamma(1)$ due to the presence of this common component.

This parameterization can be inverted and transformed to produce a reduced form MAR of the form

$$\Delta X_t = C(L)\epsilon_t \tag{12}$$

where C(0) = I and the variance-covariance matrix of ϵ_t is denoted Σ_ϵ. From (10) and (12)

$$\Gamma(L)\eta_t = C(L)\epsilon_t \tag{13}$$

and since C(0) = I the structural innovations and dynamic multipliers are given by

$$\eta_t = \Gamma(0)^{-1}\epsilon_t \tag{14}$$

$$\Gamma(L) \quad = \quad C(L)\Gamma(0) \tag{15}$$

Hence, the structural innovations and parameters can all be identified given knowledge of $\Gamma(0)$.

From (14), the covariance condition

$$\Sigma_\epsilon = \Gamma(0)\Sigma_\eta\Gamma(0)\prime \tag{16}$$

provides three restrictions which must be satisfied by the elements of $\Gamma(0)$.[15] However, there are four distinct elements to be estimated in $\Gamma(0)$. Consequently, in general there are multiple structural MAR's admissable from (10) conditional on the fourth identifying restriction imposed. Here, the maintained economic model that we seek to identify as (10) implies multiple restrictions on the long-run multiplier matrix, $\Gamma(1) = \sum_{i=0}^{\infty}\Gamma_i$, for the two disturbances which should be simultaneously satisfied.[16]

First, the aggregation of all fundamental disturbances into aggregate permanent and transitory relative price shocks, $[\eta_t^T, \eta_t^P]\prime$, implies that

$$\Gamma(1)_{21} = 0 \tag{17}$$

should hold, where the first disturbance is arbitrarily selected as the transitory relative price shock, η_t^T.[17] Second, equilibrium PPP implies that there should be no significant long-run effect for the nominal exchange rate of η_t^T or that the real exchange rate should exhibit long-run neutrality with respect to this shock:

$$\Gamma(1)_{11} = 0. \tag{18}$$

Finally, η_t^P should have equal long-run effects for the nominal exchange rate and for relative prices, also thereby satisfying a long-run neutrality criterion;

$$\Gamma(1)_{12} = \Gamma(1)_{22}. \tag{19}$$

The imposition of only one of these economic restrictions provides the fourth assumption needed to identify statistically $\Gamma(0)$ of (10); however, irrespective of which of them is imposed, all three

[15] In fact, Σ_η is normalized to be the identity matrix and the diagonal elements of $\Gamma(0)$ estimated as the impact multipliers.

[16] $\Gamma(1)$ represents the cumulative impact of each disturbance on the elements of ΔX_t. If X_t is a covariance stationary vector, then these long-run effects are zero. If X_t is non-stationary in levels and the VAR is thus specified in first differences of X_t, $\Gamma(1)$ reflects both the cumulative impact of a given disturbance on the first difference of the variable and the infinite horizon effect on the *level* of the variable.

[17] It is worth noting that the arbitrary choice of ordering for the identified disturbances involves a normalization; the point estimates are unique up to a column sign change. While in general there may be differences across choice of ordering in small samples due to sampling error, inference based on simulated standard errors computed in this paper is robust to such changes in ordering.

should be satisfied under the maintained economic model. (In fact, under the maintained economic model the imposition of any of these restrictions should induce identical empirical representations.) Inspecting estimates of the long-run multipliers, $\Gamma(1)$, and impulse response function, $\Gamma(L)$, generated by any one of the restrictions (17)-(19) therefore allows me to evaluate whether the data are consistent with the remaining long-run neutrality criteria. This method of evaluating consistency of the data with long-run neutrality restrictions embodied in (10) is discussed further in the next section.

The representation derived from the VECM parameterization that is induced by imposing (17) will be called the Baseline Model. This Model is exigent for my evaluation of PPP. The representation which is induced by imposing (18) will be called Model Ib, and the third representation which is generated by imposing (19) will be called Model Ic. These alternative representations will be studied to uncover any sensitivity of the main results to the choice of identifying assumption.

The second possible parameterization of (10), which directly imposes long-run PPP, is a reduced form cointegrated VAR in $Y_t = [\Delta(p_t - p_t^*), r_t]'$ which is estimated as

$$Y_t = B(L)Y_{t-1} + \epsilon_t \tag{20}$$

In this instance the same identification problem arises from a reduced form MAR. Now, however, there is a single, long-run restriction that suffices to identify exactly the maintained model since equilibrium PPP (and so (18) and (19)) is embodied in the specification through the inclusion of r_t as a dependent variable. Specifically, imposing that the long-run effect of one disturbance for $(p_t - p_t^*)$ is zero should retrieve (10) directly:

$$\Gamma(1)_{11} = 0. \tag{21}$$

where the first disturbance is, again, arbitrarily selected to be η_t^T. This representation will be called Model 2.[18]

For both parameterizations, the underlying structural model can be retrieved in the form of interest - as a description of the levels behaviour of exchange rates and prices following alternative disturbances - as

$$X_t = \frac{\Gamma(L)}{(1-L)}\eta_t \tag{22}$$

[18] An identical representation should be retrieved, under the maintained model, by using Δe_t as the integrated dependent variable and imposing a permanent/transitory decomposition for the nominal exchange rate. The results from this exercise are not discussed in the present paper for reasons of brevity but are available from the author upon request.

by appropriate transformation of the estimated parameters. The levels behaviour of the real exchange rate is estimated directly from the second parameterization, and can be derived by appropriate parameter transformations from the first.

B. Evaluation of Equilibrium PPP

The following methods are used to evaluate equilibrium PPP from each of the four representations of (10) described above.

The Baseline Model is estimated and the long-run multiplier matrix inspected for satisfaction of the following two overidentifying, long-run neutrality restrictions:

$$\Gamma(1)_{11} = 0 \tag{23a}$$

$$\Gamma(1)_{12} = \Gamma(1)_{22} \tag{23b}$$

Since long-run multipliers are difficult to estimate precisely, they are supplemented by information from the impulse response function of the system in levels, the $\Gamma(L)/(1-L)$.

The satisfaction of these restrictions is sufficient but not necessary for non-rejection of the long-run PPP hypothesis. If either of them is violated, equilibrium PPP can still be said to hold provided that the source of the violation accounts for a statistically insignificant fraction of the variance of the real exchange rate at long forecast horizons. Consequently, I compute forecast error-variance decompositions for relative prices, the nominal exchange rate and, by appropriate transformation of the estimated parameters, the real exchange rate. These present the percentage of the variance in the forecast errors of each variable at horizon k accounted for by the transitory and permanent relative price shocks, and are given by

$$V^j_{\eta_i}(k) = \frac{\sum_{L=0}^k \gamma(L)^2_{ij}}{\sum_{L=0}^k \gamma(L)^2_{Tj} + \sum_{L=0}^k \gamma(L)^2_{Pj}} \qquad j = 1,2; i = T, P; \tag{24}$$

where the $\gamma(L)_{ij}$ represent the estimated Lth dynamic multiplier element of $\Gamma(L)/(1-L)$ for the jth variable with respect to the ith shock. These statistics evaluate whether any deviations from equilibrium PPP observed are statistically significant sources of real exchange rate variance.

In each of these exercises, statistical inference is based on standard errors and confidence intervals derived from Monte Carlo computation of the empirical distribution of $\Gamma(L)$ under the null hypothesis that equilibrium PPP holds. This involves the assumption of independent normal errors in the reduced form VECM, $\epsilon_t \sim N(0, \Sigma_\epsilon)$, such that $S = \sum_{t=1}^T \epsilon_t \epsilon_t' \sim W_n(\Sigma_\epsilon, T)$, where W_n denotes the Wishart (matrix chi-squared) distribution. I construct a sequence of 2500 Monte Carlo draws for the reduced form error covariance matrix, Σ_ϵ, from such a distribution and use each of these to generate a draw for the reduced form VECM parameter matrix, B(L). From the parameters of each draw I can retrieve the structural model's parameters and finally compute their means, standard errors and variances where, in each case, the mean of the draws is assumed to be the "true" parameter value.[19]

Model 1b effectively identifies a permanent/transitory decomposition for the nominal exchange rate. Under the maintained model, this should exhibit long-run behaviour identical to that of the Baseline Model. I employ the techniques described above to determine whether the following neutrality restrictions are observed:

$$\Gamma(1)_{21} = 0 \tag{25a}$$

$$\Gamma(1)_{12} = \Gamma(1)_{22} \tag{25b},$$

or whether any violations of these restrictions that are uncovered account for a negligible fraction of real exchange rate variance.

Model 1c identifies the first disturbance as having equal long-run effects for e_t and $(p_t - p_t^*)$. Under the maintained economic model, only common, fundamental disturbances can cause variation in these two variables and so can be aggregated in either of the two ways described by the Baseline Model and Model 1b. Model 1c should then reproduce in some form the information embodied in the preceding Models and satisfy the single sufficient condition for long-run PPP:

$$\Gamma(1)_{11} = \Gamma(1)_{21} \tag{26}$$

Again, both disturbances are neutral for the real exchange rate at the infinite horizon when this condition holds, or if any observed violation explains a negligible portion of real exchange rate variance.

There is no unique set of overidentifying restrictions supplied by theory which can be used to evaluate Model 2's performance. Since equilibrium PPP is imposed directly on the system, evaluation of this hypothesis involves an assessment of the compatibility of the resulting *dynamics* of the system to those of economic models that are characterized by equilibrium PPP. These dynamics also should be consistent with those generated by the preceding Models. Model 2 is included for completeness; it informs theory on the dynamics that should be generated by any abstract economy that has PPP as an equilibrium condition.

[19]There are alternative methods for deriving the standard errors of structural VAR statistics. This particular method has the advantage of simplicity, at the cost of making some specific distributional assumptions.

IV. Results

I conduct all estimation over the sample period 1975:1-1991:12 for the G-7 countries Canada, France, Germany, Japan, the United Kingdom and the United States, while the Italian data were available only for the shorter sample period to 1989:12. I define the nominal exchange rate as the logarithm of the price of one U.S. dollar in domestic currency units, and define relative aggregate prices in one of two ways; first, as the logarithm of the WPI of the domestic country minus the logarithm of the WPI in the US, and second by employing the CPI in place of the WPI for each country. I use two price indices in the analysis to assess sensitivity of the results to the use of data series which reflect different concentrations of tradable goods. The CPI is commonly attributed with partial responsibility for rejections of PPP in the data due to its including a relatively large concentration of non-traded goods' prices. In the context of this paper, this amounts to an assertion that PPP fails as a result of significant changes in the equilibrium relative prices of (traded and home) goods within an aggregate commodity basket, changes which are most strongly reflected in CPI data.

The results of central interest are those generated by the permanent/transitory relative price decomposition of the Baseline Model, and I supplement these in a subsection below with sensitivity analysis based on results derived from Models 1b,1c and 2.

A. Estimation

For each country I estimate a reduced form VECM in $[\Delta e_t, \Delta(p_t - p_t^*)]'$ and a reduced form VAR in $[\Delta(p_t - p_t^*), r_t]'$, using both relative CPI's and WPI's and including four lags of endogenous variables selected by standard criteria.[20]

The coefficients on the lagged error-correction term - the real exchange rate - in the VECMs are shown in Table 1. The error-correction variable appears to play the appropriate correcting role for the nominal exchange rate; a positive deviation from PPP, due to either a positive shock to nominal rates or a negative relative price disturbance, causes a negative nominal exchange rate response in the following period. Furthermore, relative prices rise in the period following a positive PPP disturbance. There is therefore some empirical evidence for error-correcting exchange rate and price processes. However, in few cases are the coefficients on the error-correction term significant. I invert the reduced form VECM's and VAR's and estimate the structural MA Models using the identification procedures described above. [21]

[20]Specifically, the lag length was selected by examining the errors of the VAR's for evidence of serial correlation using the Q-statistic derived in Doan (1990).

[21]All computation involves solving the implied nonlinear system of equations given by (16) subject to the relevant

B. Structural Long-Run Multiplier Estimates

Estimated long-run multiplier matrices for each country and for the Baseline Model are shown in Tables 2 and 3. These provide information directly on the validity of long-run PPP in this data. In many cases the estimates are significant, in some they are highly imprecise.

For Canada, the results for the Baseline Model indicate that while the permanent relative price disturbance has an approximately equal effect on nominal exchange rates and relative prices, there is a large and significant permanent effect for nominal exchange rates of disturbances that only transitorily move relative prices. The size and significance of this nominal exchange rate effect is, moreover, independent of which price index is used to identify relative price shocks.

In the German data, the equality of relative price and nominal exchange rate responses to permanent price shocks in the Baseline Model holds only subject to sampling error but is consistent with the Canadian results. In addition, the use of alternative price indices generates a similar size for the permanent nominal exchange rate component associated with the transitory relative price shock. A remarkably similar pattern of infinite horizon responses can be seen in the Italian and UK data. The French and Japanese results warrant further comment.

The French data generate imprecise estimates for responses to the permanent relative price shock when CPI's are used. In the WPI case, the results for France are consistent with the preceding results subject to sampling error. However, for both measures of prices, the French/US nominal exchange rate exhibits a large permanent movement in response to transitory price shocks. Only for Japan is there little evidence to support the equality of relative price and nominal exchange rate long-run responses to permanent price shocks, especially when CPI's are employed, suggesting the presence of an additional source of permanent real exchange rate shocks. Again there is a significant, permanent nominal rate component not captured by permanent relative price shocks.

In general, these results show that while there are approximately equal infinite horizon effects - subject to sampling error - of permanent relative price shocks for the nominal exchange rate and relative prices as predicted by equilibrium PPP, there exist large permanent nominal exchange rate movements which occur independently of permanent relative price movements. Given the imprecision of some of the -long-run multiplier estimates, however, the next section presents the results of shorter horizon impulse response analysis.

C. Impulse Response Functions

The impulse response functions provide information regarding the behaviour of nominal ex-

linear restrictions on the elements of $C(1)\Gamma(0)$ using the GAUSS NLSYS package with the default non-linear solution algorithm and program settings.

change rates, relative price levels and the real exchange rate over a one year period following a disturbance of one standard deviation (unity) in an orthogonal innovation. For the sake of brevity, only the impulses constructed for the WPI Models are presented here, and I discuss in the text any differences observed when CPI's are employed. The results for the Baseline Model are shown in Figures 3-8.

The Canadian results indicate negligible differences when alternative price indices are employed in nominal exchange rate, relative price and real exchange rate dynamics following either permanent or transitory relative price disturbances. In both cases, the long-run effects for the levels of each variable are almost complete by the twelve month horizon. The permanent price disturbance raises relative prices and the nominal exchange rate permanently by equal amounts and has no significant real exchange rate impact at *any* lag since the nominal rate and price dynamics insignificantly differ. The transitory price disturbance accounts for no significant movement in relative prices at any lag but engenders a large, significant nominal exchange rate response at all lags. This suggests that relative prices follow a pure random walk, or have an insignificant transitory component, and that if long-run PPP fails in the Canadian data it is due to nominal exchange rate disturbances which have no effect for relative prices.

The results for France, in Figure 4, are similar in the latter respect. Transitory relative price disturbances have no significant effect for relative prices at any lag, but a large, positive and increasing effect for the nominal exchange rate. This is true when either wholesale or consumer price indices are employed. Responses to the permanent relative price disturbance contrast, however, with those of the Canadian model, and are conditional on the definition of prices used. This disturbance permanently raises both relative prices and the nominal exchange rate but here there is some evidence of a long-run PPP deviation when CPI's are used. This comprises a positive real exchange rate response that becomes significant after a six month period, although is estimated with a large standard error even at this short horizon. The real, wholesale price exchange rate is *negatively* affected by the permanent price disturbance to a twelve month horizon, but this effect subsequently dies out.[22]

The German results reflect those for Canada, as seen in Figure 5. There are no significant differences in dynamics that depend on which price index is used and in both cases there are insignificant real exchange rate effects at all lags following a permanent relative price shock. Again, the real rate is permanently raised at all horizons after a transitory price disturbance which has no significant effect for relative prices but a large, significant positive impact for the nominal exchange

[22]The sensitivity of these results to alternative price indices may be attributable to the use of an imported materials price index for France, the only available proxy for wholesale prices.

rate. The Italian data in Figure 6 and the Japanese CPI case (not shown) also reproduce these results. The use of WPI's in the Japanese case, however, generates a significant and positive real exchange rate response following a permanent price disturbance in Figure 7, reflecting the long-run multiplier estimates. Disturbances to PPP due to the transitory price shock retain the properties exhibited in the preceding cases.

The results for the UK/US system are shown in Figure 8. The dynamic effects of the transitory relative price disturbance are equivalent to those in previous cases. There are permanent components in nominal and real exchange rates that are never reflected in relative prices. Similarly, the permanent price shock generates no significant real exchange rate response at any lag, corroborating other findings of the long-run neutrality of the permanent price component for most currency prices. Once more, these results are independent of which price index is used.

This impulse response analysis based on just the first twelve months of each variable's response to alternative disturbances largely corroborates inference derived from the long-run multiplier matrices. In all cases, there are large and significant nominal rate effects of disturbances that are purely transitory for relative prices. The impulses also reveal that in no single case is the point estimate for relative price responses to this disturbance significant at *any* lag. In addition, in ten of the twelve cases studied in this sample there are no significant real exchange rate effects of permanent relative price shocks by a twelve month horizon following such a disturbance. Consequently, with the exception of parity relations involving French CPI's and Japanese WPI's, permanent disturbances to PPP are caused only by shocks that permanently move the nominal and real exchange rate but are never reflected in relative prices.

D. Forecast Error Variance Decompositions

Violations of long-run parity observed in the preceding analysis are economically relevant only if they account for a statistically significant fraction of real exchange rate variation. In other words, they can cause actual violations of equilibrium PPP if the permanent component of the real exchange rate is significant. For example, in the ten of twelve cases studied in which permanent relative price shocks display long-run neutrality with respect to the real exchange rate there would in fact be no long-run violations of PPP at all if the transitory price shock accounted for a negligible portion of real exchange rate variance.

Forecast error-variance decompositions are therefore constructed for each country and price index for the Baseline Model and the results are shown in Tables 4 and 5 for the WPI and CPI estimates respectively.

The results for Canada indicate that the permanent price disturbance accounts for a small and insignificant fraction of the total forecast error-variance in the real exchange rate although it can, naturally, explain almost all of the variance in relative prices. The fact that long-run neutrality of permanent relative price disturbances holds in the data is, therefore, irrelevant on average for unpredictable real exchange rate variation. In addition, almost all of the variance in nominal exchange rates is accounted for by the transitory price shock, so that the percentage forecast error variance of real and nominal exchange rates explained by this disturbance are very similar.

These observations suggest that while relative prices are approximately a pure random walk with respect to some permanent shock(s), real exchange rate variation is largely due to sources of permanent nominal exchange rate movements which are orthogonal to sources of variation in relative goods prices. Given these results, one would expect to observe the variation in real and nominal exchange rates to be very similar and for nominal exchange rates and relative prices to look more alike than real exchange rates and relative prices, which is in fact observed in the data.

These results are mirrored in the reports for the remaining five countries with few exceptions. In the French data, the significant equilibrium PPP violations attributable to the permanent (CPI) price disturbance observed in the impulse response functions are found to account for an insignificant fraction of real exchange rate variance (although the point estimates account for about one quarter of all violations). Consequently, the "excess" permanent nominal exchange rate component retains dominance in generating permanent PPP violations in this case. In the Italian data, while *nominal* exchange rate variation is more strongly associated with the permanent price disturbance than is typically the case, this does not generate a significant role for such disturbances in producing long-run PPP violations.

The results for Japan are the only case in which transitory price shocks are not the primary source of real exchange rate movements at all horizons. Here, permanent relative WPI shocks can account for almost one half of all real exchange rate variation at very short horizons, although they diminish in size and significance over a two year horizon. Since the impulse responses indicate the existence of permanent real exchange rate shocks due to permanent relative price disturbances, violations of long-run PPP can occur (in this case only) as a consequence of such disturbances.

Overall, the results from the Baseline Model indicate that in all twelve cases studied there are significant and permanent disturbances to PPP that are almost entirely attributable to permanent nominal exchange rate shocks and which are never reflected in relative price levels.

E. Robustness

While there is some sensitivity of the estimates to the use of alternative price indices, the main results are essentially invariant to the choice of price data, suggesting that previous criticisms of the use of consumer prices for evaluating propositions about PPP are unfounded. The Baseline Model also was estimated under alternative orderings of the Choleski matrix and statistical inference found to be robust to these orderings.

In addition, Models 1b, 1c, and 2 were estimated and equilibrium PPP evaluated in the former two cases by inspecting the identified representations for satisfaction of the over-identifying restrictions described in Section III. Model 2 is used to illustrate the empirical implications of imposing long-run PPP on a structural VAR. The full set of estimation results are presented in Betts (1994), and I document only a summary here.

The long-run multipliers derived from Model 1b, which imposes a permanent/ transitory decomposition for the nominal exchange rate, reflect the findings of the Baseline Model. In all bilateral country cases, the long-run neutrality restrictions (25a) and (25b) are violated. Relative prices respond permanently and significantly to disturbances identified to be purely transitory for the nominal exchange rate and insignificantly to the permanent nominal exchange rate disturbance. In all cases, the point estimates of long-run relative price responses to the transitory nominal exchange rate shock correspond closely to the long-run relative price response to its own permanent shock in the Baseline Model. A similar correspondence in point estimates can be observed by comparing the large and significant nominal exchange rate response to its own permanent shock in Model 1b with the nominal exchange rate response to the transitory relative price shock in the Baseline Model. These results likely reflect that the estimated long-run effects for nominal exchange rates of transitory relative price disturbances in the Baseline Model are large and significant relative to the effects of permanent relative price disturbances and so are identified as the permanent nominal exchange rate component in Model 1b.

The estimates of long-run multipliers for Model 1c almost replicate those for the Baseline Model. Since Model 1c is identified as a representation in which one of two disturbances has an equal long-run impact for relative prices and the nominal exchange rate, and the Baseline Model delivers the result that the real exchange rate exhibits such long-run neutrality with respect to permanent relative price shocks, the permanent relative price shock of the Baseline Model turns out to be exactly the "equal effect" shock of Model 1c. In every case the second shock, whose (relative) effects for the nominal exchange rate and relative goods prices are unrestricted, has a large and positive long-run impact on the nominal exchange which in half of the cases considered is significant.[23]. The relative price response to the second shock in the decomposition is small and insignificant at the infinite horizon in all cases. Thus, the second shock approximates the η_t^T identified in the Baseline Model and generates permanent disturbances to PPP.

For Model 2, the long-run multipliers are exactly statistically identified by the imposition of the restriction (21). However, the impulse response functions afford some insight for the dynamics of a system in which long-run PPP is imposed relative to the Baseline Model and to the extant theoretical literature. Those for the WPI real exchange rate are presented in Figure 9. Here, both η_t^T and η_t^P are restricted to be neutral at the infinite horizon for the real exchange rate. The preceding analysis has shown, however, violations of this condition in the responses of real and nominal rates to η_t^T, suggesting that Model 2 is misspecified. In fact, Model 2 has an impulse response function which itself suggests such misspecification.

Model 2 identifies a transitory relative price shock which has effects for the real exchange rate that are not dissipated even after a five year (sixty month) period. (In fact, relative prices for five of the countries respond *negatively* to the transitory relative WPI shock while the nominal exchange rate response is positive.) In addition, the Canadian, German and UK bilateral real rates exhibit non-neutralities for up to three years following the identified permanent relative price disturbance. When CPI's are substituted, η_t^T retains its strong persistence for PPP deviations, while the French, German, Italian and Japanese impulses display quite prolonged non-neutralities from the permanent price shocks. In all cases, the real exchange rate exceeds its long-run value following transitory price shocks but this, in contrast to the transitory "overshooting" observed in models of exchange rate dynamics a la Dornbusch, occurs with maximal impact two to eight months after the disturbance is realized.[24] Such a temporal pattern of exchange rate dynamics, and the extreme persistence of η_t^T for the real exchange rate, is difficult to reconcile with the implications of traditional exchange rate determination theory.[25]

Forecast error variance decompositions for Model 2, reported in Betts (1994), indicate that the extremely long-lived effects of the transitory price shock for real exchange rates is the only significant source of PPP disturbances at any horizon.[26] In this sense, the results for Model 2

[23]This effect is insignificant when WPI's are used for Germany, Italy and Japan, and when CPI's are employed for France, Germany and Italy

[24]This finding is consistent with the results of Eichenbaum and Evans (1993) who use multivariate, triangular Choleski decompositions to evaluate the implications of monetary policy shocks for exchange rates.

[25]These two properties of real and nominal exchange rate responses to η_t^T in Model 2 do approximate those observed in the simulated responses to monetary policy disturbances under monopolistic price setting in Beaudry and Devereux (1994); however, the responses of relative WPI's observed here are difficult to reconcile with those generated by that model, and while their monetary policy disturbance is permanent for prices here the responses are to shocks identified to be purely transitory.

[26]Two exceptions are the Canadian and German bilateral rates when WPI's are employed.

which imposes long-run PPP reflect those for the Baseline Model. The notable idiosyncracies in Model 2's impulse response functions undoubtedly reflect the imposition of a false identifying assumption in this decomposition of variance; that η_t^T evokes purely transitory disturbances to equilibrium PPP.

V. Conclusion

The results that I present in this paper indicate that in all bilateral country cases studied, there are significant departures from PPP at all horizons that are almost entirely attributable to disturbances that have permanent nominal exchange rate effects but which are never reflected in relative price levels. This is despite the fact that permanent relative price level disturbances typically are reflected one-for-one in nominal exchange rates, a long-run neutrality result that must hold in any abstract economy that delivers PPP as an equilibrium condition and which is predicted by monetary theories of the exchange rate. In fact, while the transitory relative price disturbances that generate permanent PPP deviations account for an insignificant fraction of variation in either relative consumer or wholesale price indices at any forecast horizon, they account for approximately 100% of the variance of real and nominal exchange rates. Sources of variance in relative prices and in real and nominal exchange rates are, therefore, orthogonal and there is a large, significant permanent component in real and nominal exchange rates which is orthogonal to - or excess relative to - sources of variation in relative indices of national purchasing power.

These results are robust to alternative orderings of the Choleski matrix and to alternative plausible identification schemes. In addition, whether consumer or wholesale prices are used to construct a relative price variable is irrelevant for the main results, suggesting that previous criticism of the use of consumer prices for evaluating propositions about PPP may be unfounded. Finally, when equilibrium PPP is imposed on the data representations directly this induces a decomposition with dynamics which are difficult to reconcile with the implications of extant models of exchange rate determination, and transitory shocks with extraordinarily long-lived real exchange rate effects.

These results are, of course, consistent with the empirical observation that nominal and real exchange rates are approximately equally volatile and that neither variable appears to be closely related to relative price level movements. Here, the orthogonality of exchange rate and relative price variation which is uncovered by the decompositions implies that there exist important sources of "excess" permanent nominal and real exchange rate movements that are unrelated to price fundamentals. Whether these are identifiable with some form of persistent monetary policy shock or financial market activity, as in the liquidity effects literature, or with extraneous uncertainty, cannot be determined within the context of this paper. What is clear is that the results are difficult to reconcile with "real shocks" to exchange rates deriving from changes in equilibrium relative prices of home and tradable goods within aggregate commodity baskets. Many economic models that admit permanent PPP departures predict that an important source of these departures, and of associated transitional dynamics in relative aggregate price levels, are these real price shocks. Yet here, in sample, relative aggregate prices *never* reflect permanent PPP disturbances.

Further empirical analysis is needed to uncover the source of the excess permanent nominal and real exchange rate component identified here. Finer decompositions of exchange rate variance using a richer information set would be required to achieve this, and this is left to future work.

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Table 1 - Reduced Form VECM Estimates: Error-Correction Terms

Country	Equation	Estimated Coefficient WPI Model	Estimated Coefficient CPI Model
Canada	Δe_t	-0.00	-0.01
		(-0.02)	(-1.22)
	$\Delta(p - p^*)_t$	0.02	0.00
		(2.80)**	(0.47)
France	Δe_t	-0.01	-0.03
		(-1.02)	(-1.91)
	$\Delta(p - p^*)_t$	0.002	0.002
		(0.17)	(1.03)
Germany	Δe_t	-0.01	-0.026
		(-0.99)	(-1.50)
	$\Delta(p - p^*)_t$	0.002	0.002
		(2.47)*	(1.15)
Italy	Δe_t	-0.01	-0.01
		(-0.84)	(-1.41)
	$\Delta(p - p^*)_t$	0.01	0.01
		(0.82)	(1.34)
Japan	Δe_t	-0.02	-0.01
		-(1.53)	(-1.36)
	$\Delta(p - p^*)_t$	0.000	0.01
		(0.01)	(1.72)
UK	Δe_t	-0.01	-0.02
		(-1.141)	(-1.77)
	$\Delta(p - p^*)_t$	0.01	0.002
		(1.83)	(0.69)

Table 1 presents values of the estimated error-correction terms on each of the two variables in the reduced form bivariate ECM's. These ECM's are used to identify the Baseline Model, Model 1b and Model 1c. The third column reports these estimated coefficients when the log of relative WPI's is used to measure the relative aggregate price level between the **Country** and the US. The fourth column reports results for the case in which the log of relative CPI's is employed. t-statistics are in parentheses.

Table 2 - **WPI Baseline Model Long-Run Multiplier Estimates**

Country	**Variable**	**Permanent (p-p*) Shock**	**Transitory (p-p*) Shock**
Canada	Δe_t	0.0054	0.0131
		(0.0049)	(0.0034)
	$\Delta(p - p^*)_t$	0.0056	0.0
		(0.0011)	(0.0)
France	Δe_t	0.0297	0.0383
		(0.0192)	(0.0124)
	$\Delta(p - p^*)_t$	0.0466	0.0
		(0.0141)	(0.0)
Germany	Δe_t	0.0165	0.0456
		(0.0043)	(0.0195)
	$\Delta(p - p^*)_t$	0.0087	0.0
		(0.0046)	(0.0)
Italy	Δe_t	0.0374	0.0371
		(0.0357)	(0.0108)
	$\Delta(p - p^*)_t$	0.0246	0.0
		(0.0129)	(0.0)
Japan	Δe_t	0.0458	0.0374
		(0.0506)	(0.0149)
	$\Delta(p - p^*)_t$	0.0164	0.0
		(0.0120)	(0.0)
UK	Δe_t	0.0149	0.0467
		(3.9171)	(0.0162)
	$\Delta(p - p^*)_t$	0.0169	0.0
		(0.7075)	(0.0)

The elements of **Table 2** report the response of each **Variable** to each alternative **Shock** at the infinite horizon (204 months) which is also the cumulative impact on the level of each **Variable**. Relative aggregate price levels are measured by the log of relative WPI's between the **Country** and the US. All responses reported are for the Baseline Model. Standard errors computed by Monte Carlo integration (described in the text) are in parentheses.

Table 3 - CPI Baseline Model Long-Run Multiplier Estimates

Country	Variable	Permanent (p-p*) Shock	Transitory (p-p*) Shock
Canada	Δe_t	0.0067	0.0138
		(0.0058)	(0.0033)
	$\Delta(p - p^*)_t$	0.0066	0.0
		(0.0019)	(0.0)
France	Δe_t	0.0466	0.0523
		(38.495)	(0.0188)
	$\Delta(p - p^*)_t$	0.0094	0.0
		(4.3218)	(0.0)
Germany	Δe_t	0.0221	0.0459
		(0.0468)	(0.0133)
	$\Delta(p - p^*)_t$	0.0114	0.0
		(0.0129)	(0.0)
Italy	Δe_t	0.0235	0.0461
		(0.0268)	(0.020)
	$\Delta(p - p^*)_t$	0.0201	0.0
		(0.0073)	(0.0)
Japan	Δe_t	-0.0032	0.0565
		(0.0329)	(0.0197)
	$\Delta(p - p^*)_t$	0.0090	0.0
		(0.0011)	(0.0)
UK	Δe_t	0.0190	0.0480
		(0.0356)	(0.0157)
	$\Delta(p - p^*)_t$	0.0174	0.0
		(0.0087)	(0.0)

The elements of **Table 3** report the response of each **Variable** to each alternative **Shock** at the infinite horizon (204 months) which is also the cumulative impact on the level of each **Variable**. Relative aggregate price levels are measured by the log of relative CPI's between the **Country** and the US. All responses reported are for the Baseline Model. Standard errors computed by Monte Carlo integration (described in the text) are in parentheses.

Each element in **Tables 4** and **5** reports the percentage of the **Variable's** total forecast error variance attributable to the permanent relative price level disturbance. Results reported are for the Baseline Model and relative WPI's (CPI's) measure relative aggregate price levels in **Table 4** (**Table 5**). Standard errors computed by Monte Carlo integration are in parentheses.

Table 4 - WPI Baseline Model Forecast Error-Variance Decompositions: Percentage Due To Permanent Relative Price Level Shock

Country	Variable	1 month	6 months	12 months	24 months
Canada	e_t	16.17	13.84	14.32	14.52
		(15.43)	(14.47)	(15.15)	(15.57)
	$(p - p^*)_t$	95.54	99.19	99.60	99.80
		(9.92)	(3.55)	(1.73)	(0.29)
	$(e - p + p^*)_t$	0.57	1.22	0.58	0.29
		(7.87)	(8.41)	(8.40)	(8.74)
France	e_t	43.11	48.50	44.74	41.02
		(22.98)	(21.25)	(20.41)	(20.51)
	$(p - p^*)_t$	99.17	99.77	99.89	99.95
		(10.01)	(5.68)	(3.16)	(1.41)
	$(e - p + p^*)_t$	26.22	18.25	17.14	16.68
		(20.80)	(18.36)	(18.41)	(18.88)
Germany	e_t	11.98	8.15	9.26	9.83
		(16.73)	(14.93)	(16.11)	(16.94)
	$(p - p^*)_t$	97.61	99.39	99.72	99.87
		(11.34)	(6.18)	(3.53)	(1.84)
	$(e - p + p^*)_t$	1.59	0.85	1.57	2.05
		(10.85)	(10.59)	(12.00)	(13.23)
Italy	e_t	43.59	44.58	47.19	48.97
		(21.75)	(20.61)	(20.80)	(21.45)
	$(p - p^*)_t$	90.49	97.33	98.95	99.58
		(14.28)	(6.54)	(3.01)	(1.13)
	$(e - p + p^*)_t$	7.95	7.92	8.97	9.83
		(13.79)	(14.44)	(16.05)	(17.79)
Japan	e_t	77.36	63.93	61.75	60.75
		(17.44)	(18.05)	(18.29)	(18.78)
	$(p - p^*)_t$	81.09	94.71	97.80	99.06
		(16.94)	(7.70)	(3.57)	(1.40)
	$(e - p + p^*)_t$	51.61	37.25	37.11	37.70
		(20.35)	(19.64)	(20.48)	(21.40)
UK	e_t	2.14	1.86	4.64	7.12
		(11.71)	(11.76)	(14.00)	(16.22)
	$(p - p^*)_t$	99.66	99.74	99.87	99.94
		(9.86)	(5.62)	(2.87)	(1.24)
	$(e - p + p^*)_t$	1.03	3.56	2.04	1.04
		(0.83)	(12.88)	(12.29)	(12.66)

Table 5 - CPI Baseline Model Forecast Error-Variance Decompositions: Percentage Due To Permanent Relative Price Level Shock

Country	Variable	1 month	6 months	12 months	24 months
Canada	e_t	4.24	8.36	12.66	16.00
		(9.42)	(12.87)	(14.30)	(16.37)
	$(p - p^*)_t$	93.81	97.20	98.83	99.52
		(10.14)	(4.85)	(2.12)	(0.82)
	$(e - p + p^*)_t$	2.19	1.04	0.52	0.24
		(8.18)	(7.47)	(7.98)	(9.12)
France	e_t	4.84	14.00	25.55	35.51
		(13.38)	(17.39)	(20.88)	(23.50)
	$(p - p^*)_t$	93.13	94.88	97.64	99.10
		(14.37)	(9.80)	(5.32)	(2.20)
	$(e - p + p^*)_t$	0.71	5.58	14.70	24.21
		(10.34)	(13.64)	(18.18)	(22.31)
Germany	e_t	5.93	7.64	11.49	14.99
		(12.17)	(13.17)	(15.94)	(18.77)
	$(p - p^*)_t$	96.83	98.49	99.38	99.77
		(10.17)	(5.31)	(2.53)	(0.90)
	$(e - p + p^*)_t$	0.22	1.27	2.42	3.65
		(7.95)	(9.15)	(11.51)	(14.41)
Italy	e_t	1.59	13.43	16.53	18.68
		(10.98)	(16.54)	(18.00)	(19.47)
	$(p - p^*)_t$	99.08	99.90	99.96	99.98
		(9.90)	(5.16)	(2.69)	(1.10)
	$(e - p + p^*)_t$	3.69	0.53	0.45	0.49
		(11.97)	(9.77)	(11.04)	(12.65)
Japan	e_t	0.02	0.36	0.36	0.34
		(10.56)	(11.50)	(12.25)	(13.09)
	$(p - p^*)_t$	99.40	99.56	99.80	99.91
		(10.70)	(6.58)	(3.70)	(1.62)
	$(e - p + p^*)_t$	6.25	5.11	4.83	4.63
		(15.03)	(14.62)	(14.95)	(15.46)
UK	e_t	2.88	5.15	8.07	10.75
		(11.31)	(12.85)	(14.98)	(17.18)
	$(p - p^*)_t$	95.06	97.70	99.04	99.62
		(12.21)	(6.85)	(3.35)	(1.26)
	$(e - p + p^*)_t$	1.33	0.35	0.15	0.11
		(10.32)	(10.35)	(12.63)	(12.16)