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REAL EXCHANGE RATES UNDER THE GOLD STANDARD

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ABSTRACT

Purchasing power parity is one of the most important equilibrium conditions in international macroeconomics. Empirically, it is also one of the most hotly contested. Numerous recent studies, for example, have sought to determine the validity of purchasing power parity using data from the post-Bretton-Woods float and have reached different conclusions. We assert that most such studies are flawed for two reasons. First, the post-1973 data contain, by definition, only a very limited amount of the low-frequency information relevant for examination of long-run parity. Second, the dynamic econometric techniques used to model deviations from parity are typically quite crude with respect to the modeling of low-frequency dynamics. Both deficiencies are rectified in the present paper, with dramatic results. We construct a new dataset of sixteen real exchange rates covering more than a century of the classic gold standard period, and we study deviations from parity using long-memory models that allow for subtle forms of mean reversion. For each real exchange rate, we find that parity holds in the long run.

1. Introduction

The doctrine of purchasing power parity is more than four centuries old.[1] It remains a key ingredient in modern models of exchange rate dynamics (e.g., Dornbusch, 1976) and is also widely used in policy deliberations (e.g., in the determination of target zones). The idea of purchasing power parity is simply that, when measured in the same units, the monies of different countries should command the same basket of goods. Otherwise, international arbitrage should bring about adjustments in prices, exchange rates, or both, which ultimately restore parity.

It is well known, however, that strict parity obtains only under strict conditions. Many real-world complications, including transactions costs, nontradeables, trade restrictions, exchange market intervention, taxation and changes in the terms of trade, may interfere with the workings of purchasing power parity. Moreover, the use of aggregate price indexes (with potentially different and changing market baskets) further complicates empirical investigations. Perhaps not surprisingly, then, empirical tests of purchasing power parity as a short-run proposition have failed to produce a consensus.[2]

Nevertheless, most macroeconomists (including ourselves) would agree with Dornbusch and Krugman (1976), who remark that "Under the skin of any international economist lies a deep-seated

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[1] See Bernholtz (1982).

[2] Compare, for example, McCloskey and Zecher (1976, 1984) to Frenkel (1981), Hakkio (1984) and Rush and Husted (1985).

belief in some variant of the purchasing power parity theory of the exchange rate". In particular, the hypothesis of long-run parity, that is, a tendency for the real exchange rate to revert (albeit perhaps slowly) to its parity value, is attractive. The data remain discomforting, however: deviations from parity appear highly persistent. In fact, a number of authors (e.g., Roll, 1979; Adler and Lehmann, 1983; Darby, 1983; Mussa, 1986; Diebold, 1988; Meese and Rogoff, 1988; Baillie and McMahon, 1989) have argued that real exchange rates are well approximated by martingales, so that shocks have a completely permanent effect on the levels, while changes are unpredictable. That is, they argue that there is little or no tendency for nominal exchange rates and prices to adjust in such a way as to promote purchasing power parity.

We find such behavior of the real exchange rate to be economically implausible. The real exchange rate is a relative price. Accepting the hypothesis of nonstationarity of the real exchange rate implies that it can, and will, take on an infinite range of values, given sufficient time. Such wide-ranging behavior for a relative price seems unlikely from an economic perspective, and moreover, implies unexploited opportunities for large arbitrage profits.

We shall suggest a reconciliation of the economically appealing view of purchasing power parity as a long-run equilibrium and its apparent empirical rejection. Our approach has two main parts:

(1) We study the behavior of the real exchange rates during the gold standard era, a high point of international cooperation. The gold standard era lends itself to study because it affords us long spans of data, which are precisely what is required to test hypotheses about long-run reversion of the real exchange rate to its parity value (and precisely what is lacking in studies using only data from the post-1973 float). Indeed, our shortest sample spans 74 years, while our longest spans 123 years.

(2) We model the behavior of real exchange rates using a class of long-memory models substantially more general (with respect to the low-frequency dynamics of interest) than standard time-series representations. Such generality is particularly important in the present application, because it potentially enhances our ability to discriminate slow parity reversion from non-reverting martingale behavior.

2. Historical Background

2a. Monetary History

In order to understand the subtleties of data construction and the subsequent empirical analysis, it is important to understand the monetary history of the nineteenth century. We study the United States, United Kingdom, Sweden, Belgium, France and Germany; for each country, the nineteenth century was a time of gradual movement from a silver or bimetallic monetary standard to a monometallic gold standard. Often countries were on a *de facto* monometallic standard while legally on a bimetallic standard. Table 1 lists dates of de facto gold standard adoption.[3]

The United Kingdom was on a gold standard longer than any other country. With the exception of 1798-1820, the so-called "restriction" during which the Bank of England was legally permitted to issue irredeemable paper currency, the U.K. was on a gold standard from 1750-1913.

Sweden's monetary history is also straightforward. From 1830-1872, Sweden was on a monometallic silver standard. In 1873, Sweden converted to a monometallic gold standard.

The financial histories of the other countries are more complicated. Those of Belgium and France are bound together; for most of the time Belgium was a monetary satellite of France. French coins commonly circulated in Belgium and were legal tender. In fact, from 1851-1859 Belgium issued no new coinage, relying instead on French coinage. In 1860, France and Belgium, together with Switzerland and Italy, formed the Latin Monetary Union, which made each country's coinage legal tender in the other.

For many years, Belgium and France were legally on bimetallic standards that valued silver to gold at 15.5:1. Depending on the market price, the mint ratio overvalued one metal, which then circulated as coinage. Gold was overvalued

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[3] It took some countries (for instance, France and the United States) longer to return de jure. In addition, countries frequently had been on the gold standard at times previous to those listed, but subsequently left it. After the year given in Table 1, none of the countries left the gold standard until World War I.

from 1851-1866, so Belgium and France were effectively on a gold standard. At other times silver was overvalued and an effective silver standard prevailed.[4] From 1867-1873, silver became increasingly overvalued at the mint and the countries were faced with the prospect of minting an ever increasing amount of silver. First Belgium (in 1874) and then France (in 1875) placed severe restrictions on mintage of silver and stepped up the mintage of gold. Given the small quantities of silver minted thereafter (and the very large quantities of gold), both countries had moved de facto to a gold standard.

From 1764-1857, the states that would become Germany were on a bimetallic standard that overvalued silver. In fact, in 1857 only silver circulated, and Germany formally adopted a silver standard. Following the Franco-Prussian war, however, Germany used the reparations received from France to buy gold and thus switched, both legally and effectively, from silver to gold in 1872.

The financial history of the U.S. is perhaps the most complex of all. From 1791-1861, the U.S. was on various bimetallic standards. From 1791-1834, the mint ratio overvalued silver, so the U.S. effectively was on a silver standard.[5] In 1834, the mint ratio was adjusted and gold was overvalued; the

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[4] Belgium made an abortive attempt to gain gold between 1848 and 1850 by changing the mint ratio between gold and silver. Only minor amounts of gold were minted, so we assume this effort failed.

[5] Officer (1983) contains a complete discussion of the mint ratios used during 1791-1834.

U.S. was then on an effective gold standard until 1862. In 1862, to help finance the Civil War, large amounts of inconvertible paper currency (greenbacks) were issued. The U.S. remained on a paper standard until the Resumption Act of 1873 committed the U.S. to return to a gold standard in 1879. The return was accomplished smoothly and the U.S. operated on a *de facto* gold standard, which was made *de jure* as well in 1900.

2b. Data Construction

Having briefly discussed relevant aspects of nineteenth century monetary history, let us now discuss the construction of our real exchange rate series. We study annual real exchange rates constructed from nominal exchange rates, consumer price indexes (CPI) and wholesale price indexes (WPI) for six countries: Belgium (B), France (F), Germany (G), Sweden (S), the United Kingdom (UK) and the United States (US). The starting date of each real exchange rate series is constrained by the availability of CPI and/or WPI data; for most countries, the data extend back from 1913 more than 100 years to the beginning of the nineteenth century. Table 2 details the sample periods for each of the price indexes (and the associated real exchange rates). Our final dataset contains sixteen real exchange rates, beginning at the date corresponding to WPI and/or CPI availability, and ending in 1913. The WPI rates are B/F, B/G, B/US, B/UK, F/G, F/US, F/UK, US/G, US/UK and G/UK, while the CPI rates are F/B, F/G, F/S, B/G, B/S and S/G.

With the exception of about twenty years for both the U.K. and the U.S., all of the countries in our sample were on a metallic standard (gold or silver). If two countries were on the same effective standard, their exchange rates were firmly fixed, while if they were on different standards, their exchange rates floated. We convert all the national price levels into terms of gold, consistently using the appropriate *de facto* rather than *de jure* monetary standard of the country.[6]

In order to express all price levels in terms of gold, we calculate exchange rates between the countries' circulating currency and gold, taking account of periods when countries were on a *de facto* or *de jure* silver standard and when they were on a fiat money standard.[7] This would have been easy had data been available on the exchange rate between a country on a gold standard (for example, the U.K.) and a country off gold. With the few exceptions noted below, such data apparently do not exist. Thus we were forced to assume that (1) exchange rates between a country's money and gold did not vary when the country was on a gold standard[8] and that (2) the money/gold exchange

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[6] Because the United Kingdom was on a gold standard for most of the sample period, this roughly amounts to converting all prices to pounds Sterling.

[7] A general reference for all the countries is Willis and Beckhart (1929); a specific reference for Belgium and France is Willis (1968).

[8] The assumption can be rephrased as follows: Exchange rates between countries both on a gold standard do not vary. This assumption is reasonable because any movements of the exchange rate between the gold import and gold export points would necessarily be small. Moreover, Clark (1984), Officer (1986), and Spiller and Wood (1988) indicate that variations beyond the gold points, if any, are also small. Such small changes in otherwise fixed exchange rates are dominated empirically by price level movements.

rate was determined by the market silver/gold ratio when the country was on a silver standard.[9]

We take all price levels given in terms of domestic currency from Mitchell (1980), tables I1 and I2, with the exception of the U.S., which we take from the U.S. Department of Commerce (1975) Series E40 and E51. These series were multiplied by gold/currency exchange rates to convert to price levels measured in terms of gold. The exchange rates between domestic currency and gold were calculated as follows:

Belgium: From 1832-1850, silver had a mint value of 15.5. The market value of silver in terms of gold (from the most important world market, the London market, given in Del Mar (1880)) was used to compute the exchange rate between currency and gold (that is, the silver/gold exchange rate.) From 1851-1866, gold was overvalued at the mint, so Belgium was on a *de facto* gold standard. Thus, no exchange rate was necessary to express the price levels in terms of gold. From 1867-1873, silver again circulated, so the market value of silver was again used to calculate the gold/currency exchange rate. Finally, from 1874-1913, Belgium remained on a gold standard.

France: From 1806-1874, silver had a mint ratio of 15.5:1. Silver was overvalued from 1806-1850 and 1867-1874, so it circulated. From 1806-1811 and 1814-1829, an exchange rate

[9] The assumption can be rephrased as follows: The exchange rate between a gold standard and silver standard country is determined by the market price of gold relative to silver. Using the few actual exchange rate data found, this assumption appears not to be seriously violated.

between France and the U.K. was given on page 643 of House of Representatives (1886). We used this exchange rate together with the rate between the Pound Sterling and gold to calculate a Franc/gold exchange ratio. For 1812, 1813, 1830-1859 and 1867-1874, we used the market value of silver relative to gold to compute the Franc/gold rate. During the remaining years, gold circulated, and France was on a gold standard.

Germany: From 1792-1871 Germany was either on a *de facto* or *de jure* silver standard, so the market ratio of silver to gold was used to compute the exchange rate. From 1872-1913, Germany was on a gold standard.

Sweden: From 1830-1872, Sweden was on a silver standard. In 1873 Sweden converted to a gold standard and valued its previous silver coinage at 15.813:1. Thus, from 1830-1872 we used the market price of silver in terms of gold to calculate the currency/gold exchange rate.

United Kingdom: The U.K. was off a gold standard only from 1798-1820. Andreades (1966, p. 212, p.242) reports the depreciation of the currency in terms of gold, which we used to calculate our Pound/gold series.

United States: Officer (1983) and Perkins (1978) report a series of U.S./U.K. exchange rates from 1791-1900. We used these with the Pound/gold ratio calculated for the United Kingdom to compute the Dollar/gold price. We corrected for the depreciation of the (fiat) Dollar during the Greenback era, 1862-1878, using data from Mitchell (1908).

3. Long Memory Models of Real Exchange Rate Dynamics

Before considering dynamic models for the real exchange rate, we must define it. The real exchange rate is given by

$$r_t = e_t[p_t^*/p_t] \tag{1}$$

where r_t is the time-t real exchange rate, e_t is the time-t nominal exchange rate in terms of domestic currency per unit of foreign currency, and p_t^* and p_t are time-t foreign and domestic price levels. Interpreting the price levels as measuring the foreign and domestic prices of an "average good", the real exchange rate is simply the relative price of one country's good in terms of the other country's good.[10]

The hypothesis of instantaneous purchasing power parity has an immediate interpretation in terms of real exchange rate behavior, requiring that r_t be constant. The data clearly do not satisfy this instantaneous parity hypothesis; of greater interest is whether parity holds in an appropriate long-run sense. The hypothesis of long-run parity is usefully couched in terms of the time series properties of r_t. In particular, if the effects of shocks to r_t vanish in the long run (in a sense to be defined precisely), then we say that long-run parity holds.

Now let us consider dynamic models for the real exchange

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[10] The type of "average good" depends upon the price index used. If CPI's are used, the real exchange rate is the relative price of consumption baskets. If WPI's are used, the real exchange rate is the relative price of production baskets.

rate.[11] A conventional ARIMA(p,d,q) representation is

$$\Phi(L)(1 - L)^d r_t = \Theta(L)\epsilon_t \ , \qquad \epsilon_t \sim (0,\sigma^2) \qquad (2)$$

where $\Phi(L) = 1 - \varphi_1 L - \ldots - \varphi_p L^p$, $\Theta(L) = 1 + \theta_1 L + \ldots + \theta_q L^q$, all roots of $\Phi(L)$ and $\Theta(L)$ lie outside the unit circle, and d is an integer (typically 0 or 1).

The ARIMA representation is restrictive with respect to admissible low-frequency dynamics, however, which motivates our use of the more general ARFIMA (AutoRegressive Fractionally-Integrated Moving Average) representation. In the ARFIMA representation, d is not required to be an integer.[12] The operator $(1 - L)^d$ is defined through its binomial expansion,

$$(1 - L)^d = 1 - dL + \frac{d(d-1)}{2!}L^2 - \frac{d(d-1)(d-2)}{3!}L^3 + \ldots$$

For d = 1, $(1 - L)^d$ is just the usual first-differencing filter; for non-integer d, however, it is an infinite-order lag-operator polynomial with slowly declining coefficients.[13]

The ARFIMA model belongs to the class of long-memory processes, so-named for their ability to display significant dependence between observations widely separated in time. Standard ARMA processes are short-memory, because the

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[11] In accordance with the literature, r_t should be interpreted as the natural log of the real exchange rate. In all of the empirical work reported subsequently, the log real exchange rate is the object of analysis.

[12] We present here only a cursory review as necessary for the subsequent empirical analysis. For extended discussion and references, see Diebold and Nerlove (1990).

[13] Covariance stationarity requires d < 1/2. Thus, as with integer-integrated series, one can always transform a fractionally-integrated series to covariance stationarity by taking a suitable number of integer differences.

autocorrelation (or dependence) between observations τ periods apart ($\rho_r(\tau)$) decays rapidly as τ increases. Indeed, ARMA autocorrelations decay exponentially:

$$\rho_r(\tau) \sim k^{\tau}, \qquad 0 < k < 1, \ \tau \to \infty.$$

In contrast, the defining characteristic (in the time domain) of ARFIMA processes is a slower, hyperbolic, autocorrelation decay,

$$\rho_r(\tau) \sim \tau^{2d-1}, \qquad d < 1/2, \ \tau \to \infty.$$

The intuition of long memory also emerges clearly in the frequency domain. A real exchange rate displays long memory if its spectral density, f_r, increases without limit as angular frequency tends to zero:

$$\lim_{\lambda \to 0} f_r(\lambda) = \infty.$$

For an ARFIMA series, $f_r(\lambda)$ behaves like λ^{-2d} as $\lambda \to 0$, so d parameterizes low-frequency behavior. This contrasts with the usual ARIMA model, in which the spectral density is forced to behave like λ^{-2} (corresponding to d = 1) as $\lambda \to 0$. A richer range of spectral behavior near the origin becomes possible when the integer d restriction is relaxed.

In short, the ARFIMA representation is a parsimonious low-frequency generalization of the popular ARMA class; ARMA and ARIMA representations emerge as special (and potentially restrictive) cases. Such generality in approximating Wold representations is valuable in the context of real exchange rate dynamics, because of the crucial importance of low-frequency components.

The ARFIMA model can be put into moving-average form. First, write the ARFIMA model (2) as

$$(1 - L)^d r_t = B(L)\epsilon_t,$$

where $B(L) = \Phi^{-1}(L)\ \Theta(L)$. Extracting the factor $(1 - L)$ gives

$$(1 - L)r_t = A(L)\epsilon_t = \sum_{i=0}^{\infty} a_i \epsilon_{t-i},$$

where $A(L) = (1 - L)^{1-d}B(L)$ and $a_0 = 1$. The sequence of moving-average parameters a_i, $i = 0, 1, 2, \ldots$ are called the impulse responses; they track the response of future real exchange rate changes to a unit innovation. The cumulative impulse responses,

$$c_j = \sum_{i=0}^{j} a_i \quad ,$$

$j = 0, 1, 2, \ldots$, track the response of future real exchange rate levels to the same unit innovation. Parity-reversion occurs (that is, $c_\infty = 0$) when $d < 1$. Conversely, $c_\infty = \infty$ when $d > 1$. c_∞ is finite and nonzero only in the unit-root case, $d = 1$. Perhaps more importantly, examination of the sequence of cumulative impulse response coefficients at horizons of economic interest (say, one through ten years) provides important information regarding the pattern and speed with which shocks to parity are propagated.

4. Empirical Analysis

Examination of time-series plots of the various real exchange rates makes clear the need for a class of models enabling flexible parameterization of low-frequency behavior. A representative real exchange rate, US/G, is shown in figure 1. Deviations from purchasing power parity appear pronounced and prolonged, yet there appears to be a tendency toward mean reversion.

It is unlikely, however, that standard unit root tests would have power against such slow mean reversion. In fact, a battery of standard unit-root tests applied to the real exchange rates yielded results that were mixed and hard to interpret; overall, they provided little evidence of long-run parity.[14] The possibility remains open, however, that subtle forms of reversion to parity, against which standard unit root tests may have low power, are operative.[15] We therefore proceed to estimate long-memory time series models, which provide a flexible and general parameterization of low-frequency dynamics.

The parameters of real exchange rate models allowing for fractional integration may be estimated by a variety of methods, including a two-stage semi-parametric procedure (Geweke and Porter-Hudak, 1983), approximate frequency-domain maximum likelihood (ML) (Fox and Taqqu, 1986) and exact time-domain ML (Sowell, 1990a). While the semi-parametric procedure has proved

[14] Detailed results are available upon request.

[15] Sowell (1990b) conjectures that conventional unit root tests may have low power against the long-memory alternatives entertained below. This conjecture is confirmed in the Monte Carlo analysis of Diebold and Rudebusch (1990b).

useful in a number of economic applications (e.g., Diebold and Rudebusch 1989, 1990a), it is inefficient relative to ML under correct model specification, and its distributional properties are not fully understood. Thus, ML appears to be an attractive alternative, particularly in light of the reduced computational burden afforded by approximating the Gaussian likelihood in the frequency domain.[16]

Thus, following Fox and Taqqu (1986), we exploit the fact that maximization of the Gaussian likelihood is equivalent (asymptotically) to minimization of

$$\sigma^2_T(\varsigma) = \sum_{j=1}^{T-1} [I_r(2\pi j/T)/f_r(2\pi j/T,\varsigma)] \qquad (3)$$

with respect to the ARFIMA parameter vector $\varsigma = (d, \varphi_1, \ldots, \varphi_p, \theta_1, \ldots, \theta_q)'$, where $I_r(\lambda)$ is the periodogram of r at frequency λ, and

$$f_r(\lambda,\varsigma) = |1 - e^{-i\lambda}|^{-2d} \; |B(e^{-i\lambda})|^2$$

is proportional to the spectral density of r at frequency λ. As proved by Fox and Taqqu (1986), the resulting MLE is consistent and asymptotically normal.[17]

We consider ARFIMA(p,d,q) representations for log real exchange rates, where both p and q are less than or equal to

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[16] Cheung and Diebold (1990) compare the finite-sample properties of the Fox-Taqqu (1986) and Sowell (1990a) procedures, and show that their performance is comparable in samples of the size available here.

[17] For a detailed discussion of computation and maximization of the frequency-domain likelihood for ARFIMA models, see Cheung (1990).

three.[18] Because the Akaike information criterion (AIC) and the Schwartz information criterion (SIC) have different optimality properties under different conditions, which cannot be ascertained a priori, we consider the models selected by each criterion. The models selected are generally close; in fact, they agree exactly for eleven of the sixteen real exchange rates. In cases where the models selected are not identical, the model selected by the SIC is generally more parsimonious, due to the more stringent degrees-of-freedom penalty imposed by the SIC.

ML estimates of the models selected by the AIC and SIC are reported in tables 3 and 4. The estimates were obtained by minimizing (3) using the Davidon-Fletcher-Powell algorithm. Convergence was deemed to have occurred if the change in the optimized value of (3) from one iteration to the next was less than or equal to 10^{-8}. A variety of startup parameter configurations were tried, and in each case convergence to the same vector of estimates was obtained.

The ML estimates of d are striking. For each exchange rate and for each model selected, the estimated value of d is consistent with long-run parity, that is, the unit root null is consistently rejected at conventional significance levels. For some of the exchange rates, deviations from parity appear to possess long-memory, as evidenced by d estimates significantly

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[18] To ensure covariance stationarity, and following standard practice, the models are estimated in first differences, and then converted back to levels. This involves no loss of generality, because one can always factor $(1 - L)^d$ as $(1 - L)(1 - L)^{d-1}$.

different from both 0 and -1. For other exchange rates, the deviations from parity appear to be short-memory, as evidenced by d estimates insignificantly different from 0.[19]

The estimates of the remaining autoregressive and moving-average parameters are also of interest. They are generally such that the persistence of deviations from parity implied by the model as a whole is moderately high. In particular, for those models for which the deviations from parity appear to be short-memory (that is, the estimated value of d is insignificantly different from 0), the configuration of the remaining autoregressive and moving-average parameters nevertheless implies substantial shock persistence. Consider, for example, the US/G rate, which was graphed earlier. The models selected by the AIC and SIC are identical and insignificantly different from an AR(1) in levels. Estimation of the AR(1) model yields a parameter of .81, which implies that the half-life of a shock that moves the real exchange rate away from its parity value is approximately three years.

Graphical analysis of the cumulative impulse response functions enables direct assessment of the speed and pattern with which shocks to purchasing power parity are transmitted. In figure 2 we present cumulative impulse response functions for the

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[19] Note that, by estimating fractionally-integrated models, we can directly assess the amount of uncertainty associated with low-frequency variation in real exchange rates by examining the confidence bands for the estimate of d. This stands in marked contrast to the common practice of <u>conditioning</u> upon an <u>assumption</u> of d = 0 or d = 1 (typically after some pretesting.)

models selected by the AIC for each of the 16 currencies. Reversion to parity is evident in the their eventual decay toward zero; in fact, most of the impulse responses are monotone decreasing. The half-life of a shock to parity, averaged across the currencies, is 2.8 years.

5. Concluding Remarks

An emerging literature, of which our paper is a part, poses a serious challenge to the view that deviations from purchasing power parity are well approximated by martingales. In two papers that foreshadow much subsequent work, Frankel (1986, 1989) argues that the martingale hypothesis for the US/UK real exchange rate can be rejected with a sufficiently large sample. Lothian (1990), using a long sample of Japanese real exchange rates, agrees. Additionally, Huizinga (1987), Kaminsky (1987) and Glenn (1989) find some evidence of reversion to parity using variance-ratio tests, as do Abuaf and Jorion (1990), who make use of multivariate techniques.

Our paper represents a culmination of the emerging literature, building on Hakkio's (1986) conjecture that the small samples and naive techniques frequently employed in studies of real exchange rate dynamics might produce low power against alternatives of slow parity reversion. We make use of (1) long spans of data, (2) flexible time-series representations, and (3) a variety of currencies, with dramatic results. We find that purchasing power parity holds in the long run for each of the currencies studied, and that the typical half-life of a shock to parity is approximately three years.

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Table 1
Year of Adoption of the Gold Standard

Belgium	1874
France	1875
Germany	1872
Sweden	1873
United Kingdom	1821
United States	1879

Table 2
Price Index Sample Periods

Country	WPI	CPI
Belgium	1832 - 1913	1835 - 1913
France	1806 - 1913	1840 - 1913
Germany	1792 - 1913	1820 - 1913
Sweden	-----	1830 - 1913
U.K.	1798 - 1913	-----
U.S.	1791 - 1913	-----

Table 3
Parameter Estimates of Models Selected by the Akaike Criterion

	d	ϕ_1	ϕ_2	ϕ_3	θ_1	θ_2	θ_3
<u>WPI Rates</u>							
B/F	.49	----	----	----	0.37	----	----
	(.12)				(.14)		
B/G	-.20	0.76	----	----	0.47	----	----
	(.27)	(.16)			(.13)		
B/US	-.06	0.79	----	----	----	----	----
	(.18)	(.11)					
B/UK	-.30	0.86	----	----	----	----	----
	(.17)	(.10)					
F/G	.66	----	----	----	0.74	0.18	-0.21
	(.20)				(.21)	(.24)	(.14)
F/US	-.46	1.57	-0.63	----	----	----	----
	(.27)	(.27)	(.25)				
F/UK	-.41	0.91	----	----	0.53	----	----
	(.14)	(.06)			(.10)		
US/G	-.05	0.84	----	----	----	----	----
	(.18)	(.11)					
US/UK	-.38	1.31	-0.40	----	----	----	----
	(.38)	(.41)	(.26)				
G/UK	.65	0.46	-0.28	----	----	----	----
	(.17)	(.19)	(.05)				
<u>CPI Rates</u>							
F/B	.02	0.82	----	----	----	----	----
	(.34)	(.22)					
F/G	.11	----	----	----	0.96	0.52	----
	(.12)				(.12)	(.16)	
F/S	.30	----	----	----	1.03	0.71	0.28
	(.15)				(.17)	(.21)	(.14)
B/G	-.13	----	----	----	0.98	0.42	----
	(.13)				(.15)	(.18)	
B/S	.26	----	----	----	0.35	----	----
	(.04)				(.11)		
S/G	-.12	----	----	----	1.01	0.59	----
	(.12)				(.11)	(.10)	

Note to table 3: Asymptotic standard errors appear in parentheses.

Table 4
Parameter Estimates of Models Selected by the Schwartz Criterion

	d	ϕ_1	ϕ_2	ϕ_3	θ_1	θ_2	θ_3
<u>WPI Rates</u>							
B/F *	.49	----	----	----	0.37	----	----
	(.12)				(.14)		
B/G *	-.20	0.76	----	----	0.47	----	----
	(.27)	(.16)			(.13)		
B/US *	-.06	0.79	----	----	----	----	----
	(.18)	(.11)					
B/UK *	-.30	0.86	----	----	----	----	----
	(.17)	(.10)					
F/G	.35	----	----	----	0.99	0.51	----
	(.09)				(.11)	(.11)	
F/US *	-.46	1.57	-0.63	----	----	----	----
	(.27)	(.27)	(.25)				
F/UK	.50	----	----	----	0.56	----	----
	(.10)				(.10)		
US/G *	-.05	0.84	----	----	----	----	----
	(.18)	(.11)					
US/UK *	-.38	1.31	-0.40	----	----	----	----
	(.38)	(.41)	(.26)				
G/UK	.57	----	----	----	0.50	----	----
	(.09)				(.08)		
<u>CPI Rates</u>							
F/B	.80	----	----	----	----	----	----
	(.12)	----					
F/G *	.11	----	----	----	0.96	0.52	----
	(.12)				(.12)	(.16)	
F/S	.69	0.65	-0.35	----	----	----	----
	(.12)	(.11)	(.04)				
B/G *	-.13	----	----	----	0.98	0.42	----
	(.13)				(.15)	(.18)	
B/S *	.26	----	----	----	0.35	----	----
	(.04)				(.11)		
S/G *	-.12	----	----	----	1.01	0.59	----
	(.12)				(.11)	(.10)	

Note to table 4: Asymptotic standard errors appear in parentheses. An asterisk indicates agreement between the SIC-selected and AIC-selected models.

FIGURE 1

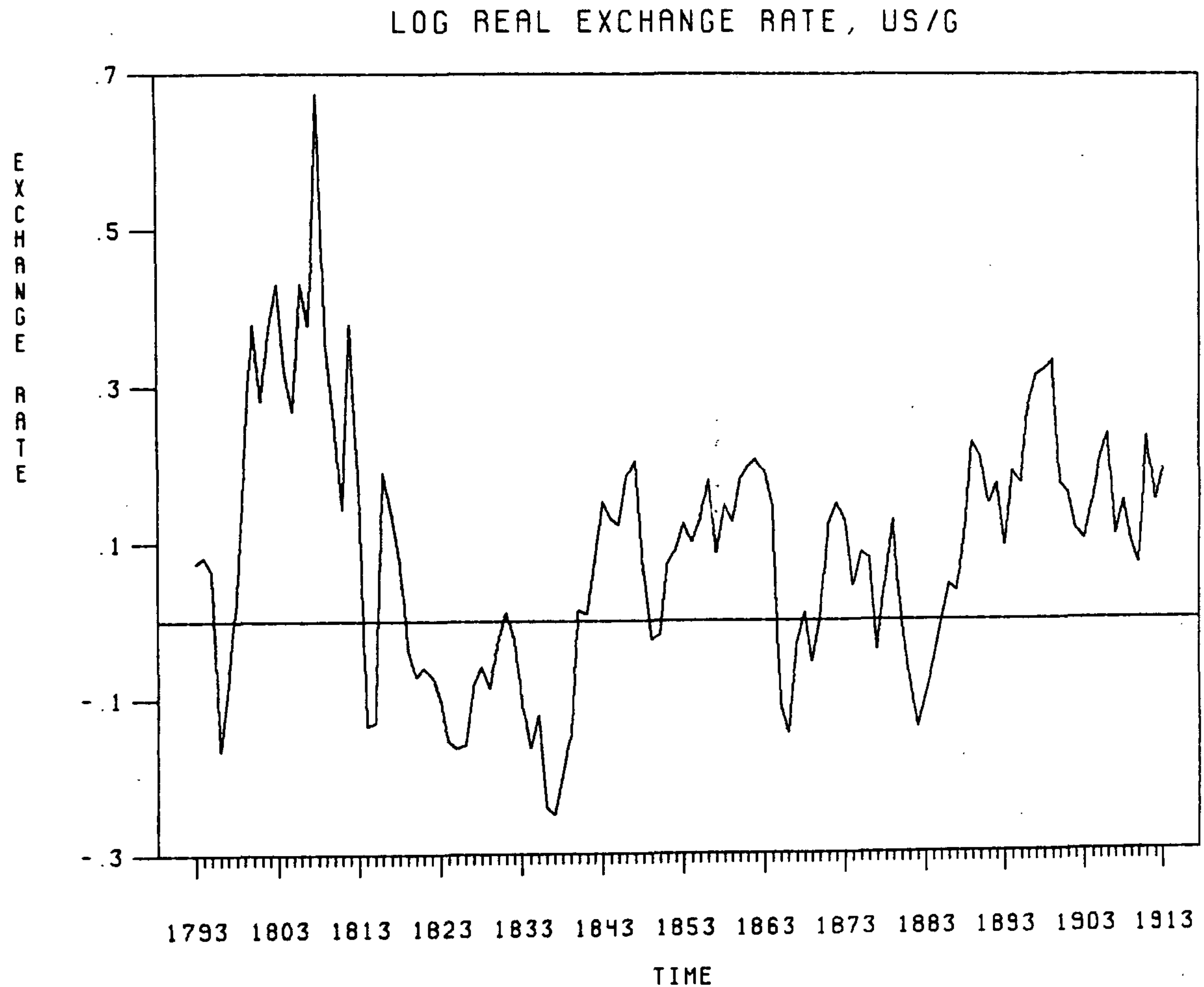
LOG REAL EXCHANGE RATE, US/G
EXCHANGE RATE
.7
.5
.3
.1
-.1
-.3
1793
1803
1813
1823
1833
1843
1853
1863
1873
1883
1893
1903
1913
TIME

FIGURE 2
CUMULATIVE IMPULSE RESPONSE FUNCTIONS

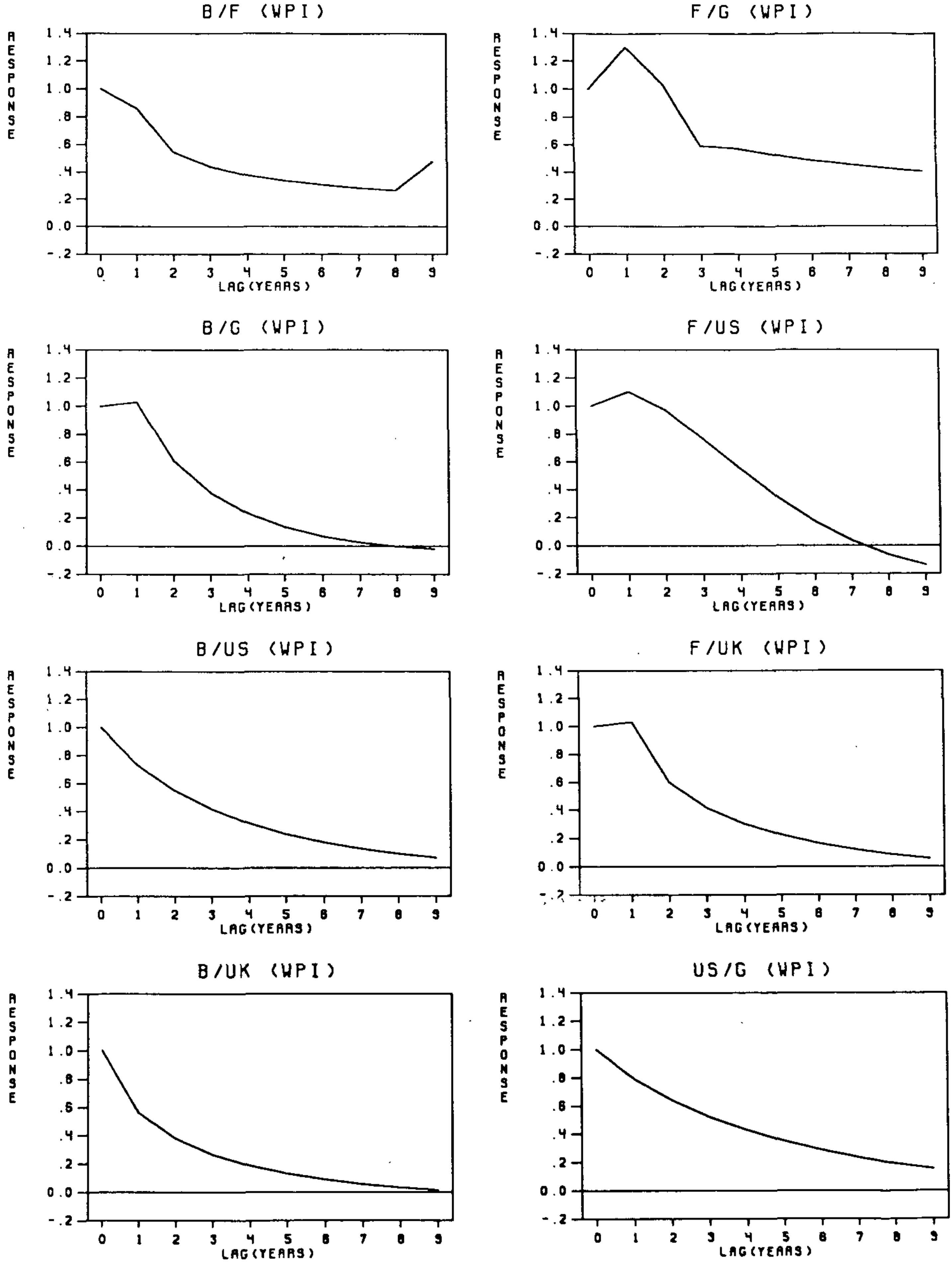

FIGURE 2 (CONTINUED)
CUMULATIVE IMPULSE RESPONSE FUNCTIONS