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Abstract:

Long half-lives of real exchange rates are often used as evidence against monetary sticky price models. In this study we show how exchange rate regimes alter the long-run dynamics and half-life of the real exchange rate, and we recast the classic defense of such models by Mussa (1986) from an argument based on short-run volatility to one based on long-run dynamics. The first key result is that the extremely persistent real exchange rate found commonly in post Bretton Woods data does not apply to the preceding fixed exchange rate period in our sample, where the half-live was roughly half as large. This result suggests a reinterpretation of Mussa's original finding, indicating that up to two thirds of the rise in variance of the real exchange rate in the recent floating rate period is actually due to the rise in persistence of the response to shocks, rather than due to a rise in the variance of shocks themselves. This result also suggests a way to resolve the "PPP puzzle," reconciling real exchange rate persistence with volatility. The second key result explains the rise in persistence over time by identifying underlying shocks using a panel VECM model. Shocks to the nominal exchange rate induce more persistent real exchange rate responses compared to price shocks, and these shocks became more prevalent under a flexible exchange rate regime.

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I. Introduction

There is a long-standing and ongoing debate over what type of model best explains the behavior of the real exchange rate, including its volatility, persistence, and comovement with the nominal exchange rate. On one hand, sticky price models argue that shocks to the nominal exchange rate are passed on to the real exchange rate because of sticky nominal prices. On the other hand, advocates of real models of the real exchange rate – "real-real models" -- argue that movements in the nominal exchange rate primarily reflect effects passed on from shocks to the real exchange rate. In his 1986 Carnegie Rochester paper, Michael Mussa offered a highly influential critique of real-real exchange rate models by observing that the volatility of the real exchange rate is higher under a flexible nominal exchange rate regime than under a fixed exchange rate regime, such as the Bretton Woods system. This finding often is used to support a sticky price story, as it is thought that the more flexible nominal exchange rates prevailing in the post-Bretton Woods period permitted a rise in the volatility of nominal shocks.

Since the time that Mussa wrote his paper, the study of real exchange rates has shifted from examining exchange rate volatility -- what Mussa referred to as "short-term fluctuations" -- toward the analysis of the long-run behavior of the real exchange rate.[1] In particular, most recent research has focused on asking if deviations in the real exchange rate from its purchasing power parity (PPP) equilibrium level tend to disappear in the long run, and how long it takes the real exchange rate to return to its long run level. Many of these papers have criticized sticky price models for their inability to explain the evident persistence of real exchange rate deviations from PPP.[2] As a recent example of this debate, Steinsson (2008) argued that the hump-shaped dynamics and long half-life of the real exchange rate in the flexible exchange rate period cannot be explained by a sticky price model in terms of nominal shocks.

[1] In his analysis Mussa (1986) worked with data in first differences (at the quarterly frequency) and consistently referred to his results as "short-term" fluctuations.

[2] For a summary, see Rogoff (1996). For a demonstration of how sticky price models have worked to overcome this criticism, see Carvalho and Necchio (2010).

In this paper we counter this argument with an updated version of the Mussa critique of real-real exchange rate models. We ask whether recent findings regarding persistent dynamics in the literature studying the standard post-Bretton Woods dataset apply also to the preceding Bretton Woods period of generally fixed exchange rates. In particular, we analyze whether the dynamics or half-life of the real exchange rate are affected by a change in the nominal exchange rate regime. If the real exchange rate is driven by real shocks, then the Mussa critique should apply and a change in the nominal exchange rate regime should not affect the dynamics of the real exchange rate. But if the change in the nominal regime does affect exchange rate dynamics, then sticky price models potentially may provide an alternative explanation. To shed further light on whether they do, we also decompose real exchange rate changes into underlying movements associated with shocks to the nominal exchange rate and to relative national prices.[3]

Our methodology takes advantage of recent advances in panel econometrics and applies them to estimating the dynamics properties of the real exchange rate. Specifically, we adapt the method of Pesaran (2006) to estimate an autoregression of the real exchange rate over the Bretton Woods and post-Bretton Woods period for a panel of 20 industrialized countries using the dollar as a common numeraire. This methodology controls for the contemporaneous correlation across country pairs in the panel. We also estimate a two-equation vector error correction model (VECM) that decomposes the real exchange rate into its nominal exchange rates and the relative price components, and use this framework to identify distinct shocks to these two components. [4]

[3] Mussa (1996) reported a variety of statistics for the nominal exchange rate and price indices as well as for the real exchange rate, including the variance, covariance, mean, and serial correlation. Interestingly, in his list of empirical observations, he notes rise in persistence under flexible exchange rates: "Short-term changes in nominal exchange rates and in real exchange rates show substantial persistence during subperiods when the nominal exchange rate is floating." We take it as reassuring that his observations coincide with our claim of greater persistence during the post-Bretton woods period, even if his main focus was on short-run fluctuations rather than on long-run convergence and half-lives of the real exchange rate to PPP equilibrium.

[4] Throughout the paper we analyze the real exchange rate in levels form rather than in first differences. Unit root tests presented in the paper support the long-run stationarity of the real exchange rate, and our VECM methodology makes use of the cointegrating relationship between the nominal exchange rate and relative prices in levels.

Our paper contributes to the empirical literature on real exchange rates with two main findings. The first contribution is the finding that the dynamic properties of the real exchange rate differ between the Bretton Woods and post-Bretton Woods periods. The half-life we estimate for the fixed exchange rate period, roughly two years, is about half as long as that for the flexible exchange rate period of about four years. Further, the hump shape identified in the real exchange rate dynamics during the flexible exchange rate period is not present during the fixed exchange rate period. The finding is surprising, as theories going back to Friedman (1953) maintain that a flexible exchange rate should be useful as an alternative adjustment mechanism of relative prices when nominal prices are not free to adjust.[5] This would suggest that the imposition of a fixed exchange rate should raise the half-life of the real exchange rate rather than lower it. The finding that the nominal exchange rate regime affects the half-live of the real exchange rate offers a new stylized fact that models of the real exchange rate should account for.

An additional implication of this finding is that it suggests a reinterpretation of the original finding of Mussa (1986) regarding short-run volatility. Our estimates imply that about two thirds of the increase in the variance of the real exchange rate under flexible exchange rates was due to a rise in the persistence of the response of the real exchange rate to shocks (the "intrinsic component," to use the terminology of Obstfeld and Stockman, 1985), rather than a rise in the variance of exogenous shocks themselves (the "extrinsic component"). This finding also suggests an approach to resolving the "PPP puzzle." The real exchange rate literature often views the stylized facts of high volatility and high persistence of real exchange rate deviations as two observations that are difficult to explain together. Our analysis suggests a role for dynamics in explaining both facts, and suggests that research focus on searching for a model that explains persistence as a means of explaining volatility.

[5] Recent theoretical contributions show that Friedman's conclusions are conditional upon assumptions such as exports invoiced and sticky in the currency of the exporter. See Berka, Devereux, and Engel (2012) for a recent discussion of Friedman's claims.

The second contribution of the paper is to investigate why the persistence of real exchange rate deviations has increased in the transition to the post-Bretton Woods period, using a two-equation VECM estimation to identify underlying shocks to the nominal exchange rate and to prices. Conditioning on these shocks turns out to be informative. First, simulation of impulse responses shows that the half-life of real exchange rate deviations due to nominal exchange rate shocks is four to six times larger than that due to price shocks, and this applies similarly to both exchange rate regime periods. Second, variance decompositions show, not surprisingly, that the share of real exchange rate fluctuations due to nominal exchange rate shocks rises during the flexible rate period. The rise in volatility of nominal exchange rate shocks that imply greater persistence seems to be the reason for the rise in overall real exchange rate persistence. When we condition by shock, most of the difference in half-lives of the real exchange rate between the two regime periods disappears.

In sum, our findings reaffirm the broad conclusion of Mussa, in that the change in the behavior of the real exchange rate during the fixed and flexible exchange rate regime periods can be attributed to a change in the incidence of shocks to the nominal exchange rate. We would emphasize that the basis of this reaffirmation differs in two respects from Mussa's analysis. First, we study dynamics, rather than the volatility, of exchange rate changes. Secondly, while the past literature has spoken only vaguely about the possible role of an increase in exchange rate shocks since the demise of the Bretton Woods system, we employ time series tools to identify such shocks econometrically in order to evaluate this claim.

This paper is related to the very large literature estimating half-lives of the real exchange rate, though typically only using post-Bretton Woods data.[6] A number of studies use long horizon data that encompasses the Bretton Woods period and earlier, but they generally do not compare half lives across regimes, nor do they employ panel estimation methodology.[7] Only a

[6] See Imbs et al (2005) as a prominent example and for discussion.

[7] For example, Abuaf and Jorion (1990) find average half lives of 3.3 years for bilateral real exchange rates between the United States and eight countries for 1900 to 1972. Murray and Papell (2002)) find an average half life of 3.6

very small number of papers in this literature have estimated half-lives for different exchange rate regimes, but they are not motivated by the question of evaluating real versus monetary models, and they do not use panel methodology (see Taylor, (2002); and Sarno and Valente, 2006).[8] Taylor (2002) is most similar in spirit, in that it tries to explain the rise in volatility of the real exchange rate under flexible exchange rate regimes, and decomposes this real exchange rate volatility into persistence and size-of-shock components.

Our methodology for decomposing real exchange rate changes into their underlying components is also closely related to Cheung, Lai, and Bergman (2004), who first showed that the response of the real exchange rate to a nominal exchange rate shock was more persistent than that to a price shock using a VECM. However, they are interested only in the flexible exchange rate period, and their empirical analysis examined data only for 5 currency pairs. Further, they estimate separate VECMs for each individual currency pair, while we estimate a panel VECM by adapting Pesaran's (2006) CCEP estimator to a VECM setting. Bergin, Glick, and Wu (forthcoming) also applied this estimation strategy, but with the objective of comparing the dynamics of disaggregated goods-level data to that of aggregates, rather than comparing across exchange rate regimes.

The paper is organized as follows. The data and preliminary analysis involving stationarity tests are presented in Section II. The main empirical results are in Sections III and IV, where section III estimates half-lives from single-equation autoregressions, and section IV

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for six countries over the period 1900 to 1996. Frankel (1986) uses a 116-year-long dataset for the dollar/pound real exchange rate and reports a half-life of 4.6 years. Lothian and Taylor (1996) find a half live of 4.7 years for the dollar-pound rate using two centuries of data. In contrast to our analysis, none of these papers compare the behavior of the real exchange rate across periods.

[8] Our results differ from Taylor (2002), who found only a modest increase in the half-lives of countries between the Bretton Woods period and the floating period that followed. In Taylor's estimates pooling across countries, the half-life rose modestly from 1.5 to 2.1 years; when he estimated separately by country, the mean rose only from 2.4 to 2.6 years. For a discussion, see our robustness checks at the end of section 3, where we apply our methodology to Taylor's data. Our methodology very closely reproduces his pooled result, showing that our main result is not due to our panel methodology. Our longer half-life for the floating period is due primarily to the focus on developed economies and the longer sample period we have for the floating period. Sarno and Valente (2006) differ from our main result in finding *faster* convergence under fixed exchange rate regimes. However, this difference appears to arise from their decision to aggregate the Bretton Woods period together with the Gold Standard as two examples of fixed exchange rate regimes. Again, see our discussion at the end of section 3 for details.

explains this result by estimating a VECM and conditioning on shock. Section V concludes the paper.

II. Data and Preliminary Analysis

The dataset consists of bilateral nominal exchange rates with the U.S. dollar as the numeraire and consumer prices indices, for 20 industrialized countries, taken from the International Financial Statistics.[9] The benchmark sample is annual in frequency and covers the period 1949 to 2010. We also show the robustness of results using monthly data, though this limits the data range to starting in 1957.

Define the real exchange rate, $q_{j,t}$ as the relative price level between country *j* and the base country (here the U. S. dollar) in period *t*, computed as $q_{j,t} = e_{j,t} + p_{j,t}$, where $e_{j,t}$ is the nominal exchange rate (currency *j* per U.S. dollar), and $p_{j,t} = p_{US,t} - p^*_{j,t}$ is the log difference between the price indices in the United States and country *j*, both in local currency units, and all variables are expressed in logs.[10]

As preparation for the main analysis later, we first establish that international relative prices are stationary. We apply the cross-sectionally augmented Dickey-Fuller (CADF) test provided by Pesaran (2007). Consider the following regression:

$$\Delta q_{j,t} = a_j + b_j(q_{j,t-1}) + c_j(\overline{q}_{t-1}) + d_j(\Delta \overline{q}_t) + \varepsilon_{j,t} \qquad j = 1,...,N, \text{ and } t = 1,...,T \tag{1}$$

[9] The full list of 20 countries is: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, and the United Kingdom. For the monthly sample, sufficient data are available only for 17 countries, with Australia, Denmark, and New Zealand being excluded. Other studies of PPP using aggregate data during the post-Bretton Woods period use similar country samples (e.g. Papell 2002). The nominal exchange rate data are defined as yearly averages.

[10] This specification assumes that $p_{US,t}$, $p^*_{j,t}$ share similar convergence speeds, a property that has been found to be consistent with the data; see Cheung et al (2005).

where $\overline{q}_t = \sum_{j=1}^{N} q_{j,t}$ is the cross-section mean of $q_{j,t}$ across the *N* country pairs and $\Delta\overline{q}_t = \overline{q}_t - \overline{q}_{t-1}$.

The purpose for augmenting the cross-section mean in the above equation is to control for contemporaneous correlation among $\varepsilon_{j,t}$. The null hypothesis can be expressed as $H_0 : b_{ij} = 0$ for all *ij* against the alternative hypothesis $H_1 : b_j < 0$ for some *j*. The test statistic provided by Pesaran (2007) is given by

$$CIPS(N,T) = (1/N)\sum_{j=1}^{N} t_j(N,T)$$

where $t_j(N,T)$ is the t statistic of b_j in equation (1).

The top panel of Table 1 indicates rejection of nonstationarity at the 10% level for all subsamples of the annual data: the Bretton Woods sample, 1949-1971; the post-Bretton Woods sample, 1973-2010; and the whole sample combined.[11] Further, the longer data length afforded by the whole sample rejects nonstationarity even at the 1% significance level. We will use the annual data as our main dataset. Using higher frequency monthly data further improves power to reject nonstationarity for the post-Bretton Woods sample. But it reduces power for the Bretton Woods sample since the monthly data begin in 1957, significantly shortening the time span. The bottom panel of the table shows that the shorter monthly sample is not able to reject nonstationarity for the Bretton Woods period.

III. Conditional on Regime: Single Equation Autoregressions

We first estimate the speed of convergence toward stationarity by estimating an autoregressive model of real exchange rates with panel data. To control for contemporaneous correlation of residuals, we apply the pooled common correlated effects (CCEP) regressor of

[11] We omit the transition year 1972 from the sample.

Pesaran (2006) to estimate the autoregressive coefficients of real exchange rates. More specifically, we estimate the equation:

$$q_{j,t} = c_j + \sum_{m=1}^{M} \rho_{j,m}(q_{j,t-m}) + \varepsilon_{j,t} \quad (2)$$

augmented with cross-section means of the left-hand and right-hand variables ($\overline{q}_t$, $\overline{q}_{t-1},...\overline{q}_{t-M}$).[12] To control for potential bias in CCEP estimators, we employ the standard double bootstrap procedure of Kilian (1998) with 1000 replications to obtain bias-adjusted estimates.[13] Specification tests based on AIC suggest one lag is optimal for annual data (M = 1), but we also report results for two lags (M = 2) to demonstrate that our conclusions are robust.[14]

Table 2 reports half-lives of the real exchange rate for an autoregession with one lag, computed on the basis of simulated impulse responses.[15] The half-life estimated for the fixed exchange rate Bretton Woods period is 2.27 years (with a 5%-95% band of 1.43 to 3.39 years). That for the more flexible exchange rate post-Bretton Woods period is 4.31 years (with a band of 3.44 to 5.24). The latter number is within the "consensus" range in the empirical literature of between 3 to 5 years.[16] We find that the half-life under fixed exchange rates was distinctly smaller than under the flexible rates, being just about half the size. A similar result obtains when we consider two lags in the autoregression: with half-lives of 2.75 and 4.27, respectively.

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[12] STATA code to conduct CCEP estimations used throughout the paper are available upon request.

[13] See the appendix of Bergin, Glick, and Wu (forthcoming) for a Monte-Carlo study of the bias of the CCEP estimator when applied to models with a lagged dependent variable. In our bootstrap methodology we do block resampling of residuals (recentered by currency pair) with replacement and initialize with actual data. We use block sizes of two periods; the results are not affected by using different block sizes.

[14] We applied the AIC in two ways: (i) to the full system of residuals from the panel estimation, and (ii) to the residuals of individual AR equations estimated for each currency pair individually. The first approach indicated an optimal lag length of 1 year for both the Bretton Woods and post Bretton Woods periods, The second approach yielded a median estimate across countries of 1 year for the post Bretton Woods period and of 2 years for the Bretton Woods period. We chose to use a common length of 1 year for both periods in our benchmark case using annual data.

[15] The half-life is computed as the time it takes for the impulse responses to a unit shock to equal 0.5, as defined in Steinsson (2008). We identify the first period, t_1, where the impulse response $f(t)$ falls from a value above 0.5 to a value below 0.5 in the subsequent period, t_1+1. We interpolate the fraction of a period after t_1, where the impulse response function reaches a value of 0.5 by adding $(f(t_1) - 0.5))/ (f(t_1) - f(t_1+1))$.

[16] See Rogoff (1996) and Imbs et al (2005). Our estimated half-life for the post-Bretton Woods period is larger than that found in Bergin, Glick, and Wu (forthcoming). As explained in that paper, the faster convergence speed is attributable to the use of product-level data that restricts the sample to the period 1990-2007.

Figure 1 plots the impulse responses associated with these half-lives. Observe in the lower panel that when two lags are included in the autoregression a hump-shaped response can be detected for the post-Bretton Woods data, as observed in Cheung, Lai, and Bergman (2005) and Steinsson (2008). However, the upper panel reveals no hump-shaped response in the Bretton Woods period. If the hump-shaped dynamics are a reason for the high degree of persistence in recent real exchange rate data, as Steinsson conjectures, then the lack of this hump may help explain the lower degree of persistence in the Bretton Woods period.

As an additional robustness check, we report results for data at a higher, monthly frequency. Since the monthly data start in 1957 instead of in 1948 for the annual data, the length of our Bretton-Woods sample is severely restricted.[17] As noted in Section II, the shorter sample for monthly data is unable to reject nonstationarity of the real exchange rate for the Bretton Woods period. Because estimating equation (2) is meaningful only when q is stationary, we chose to use annual data for our main results. But given the success of the annual data in rejecting nonstationarity, we conjecture that the failure to reject nonstationarity in the monthly frequency is due to the power loss from a much shorter time span. We nonetheless report results for monthly data (with 5 lags) to demonstrate the robustness of our claims.[18] The bottom panel of Table 2 reiterates the conclusion from annual data: the half-life is about half as long during the Bretton Woods period as afterward. Impulse responses in Figure 2 show a pronounced hump-shaped response for post-Bretton Woods data, but a clear lack of any hump during Bretton Woods.

This shorter half-life during the Bretton Woods period is surprising, as theories dating back to Friedman (1953) posit that a flexible exchange rate should be useful as an alternative adjustment mechanism of relative prices when nominal prices are not free to adjust. This suggests that the imposition of a fixed exchange rate should raise the half-life of the real

[17] Using quarterly data during the Bretton Woods period suffer even more from this problem, as the start date is later than with monthly data, implying even fewer observations.
[18] A lag length of 5 was chosen by using the AIC criterion to find the optimal lag length for each country and then taking the median value across countries.

exchange rate rather than lower it. It is also surprising in the context of nonlinear models like Taylor, Peel, and Sarno (2001) and Sarno and Valente (2006), since they predict that large shocks should be corrected more quickly, and nominal exchange rate shocks tend to be larger in magnitude than aggregate price shocks. As we show below, the effect of a flexible exchange rate as an adjustment mechanism is outweighed by its effect as an additional source of shocks.

The finding of greater persistence after Bretton Woods ended also provides a new interpretation for the long-standing finding regarding the rise in the variance of the real exchange rate in this period. The variance of the real exchange rate can be decomposed into the extrinsic and the intrinsic components, to borrow the terminology from Obstfeld and Stockman (1985). The former represents the exogenous shock to the autoregression in equation (2), and the latter is the endogenous propagation characterized by the dynamic parameter ρ in that equation. If one transforms equation (2), in the case of a single lag, by subtracting the mean of the real exchange rate from each side, squaring, and taking expectations, one derives:

$$\text{var}\left(q_{j,t}\right) = \left(\frac{1}{1-\rho_j^2}\right) \cdot \text{var}\left(\varepsilon_{j,t}\right). \tag{3}$$

This formula provides a clean decomposition of the variance of the real exchange rate, where $\left(1/(1-\rho_j^2)\right)$ is the contribution of the intrinsic component, and $\text{var}\left(\varepsilon_{j,t}\right)$ is the extrinsic component. In our dataset the variance of the real exchange rate, $\text{var}\left(q_{j,t}\right)$, doubled in the post-Bretton Woods sample compared to the Bretton woods sample, with the percent change averaging 102.45% across countries with annual data. Using the estimates of the autoregression from Table 2, we compute that the term $\left(1/(1-\rho_j^2)\right)$ increased by 66%. (See Table 3 for details.) Together, this indicates that nearly two thirds of the rise in real exchange rate variance under flexible exchange rates in our data was due, not to a rise in the extrinsic shocks to the real exchange rate, but instead to the rise in persistence under the flexible rate regime. This underscores the importance of understanding the reason for the change in dynamics between the

two regimes. Not only is the change in dynamics due to the change in exchange rate regime a potential new stylized fact of interest, but it also is important for understanding the rise in volatility itself, which long has been identified in the literature as a key stylized fact.[19]

This result is robust when applied also to the other estimations and samples discussed previously in the paper, although the derivation of the decomposition is more complicated than for the simple autoregression with one lag. Applying the decomposition to the autoregression estimated with two lags, i.e. $M = 2$, equation (2) implies the following (see the appendix for explanation):

$$\text{var}\left(q_{j,t}\right) = \Psi \cdot \text{var}\left(\varepsilon_{j,t}\right), \text{ where } \Psi = \frac{1}{1-\rho_{j,1}^2-\rho_{j,2}^2-\left(2\rho_{j,1}^2\rho_{j,2}/\left(1-\rho_{j,2}\right)\right)}. \quad (4)$$

Substituting in the autoregressive parameter values from the AR(2) estimation, the result is that $\Psi = 2.68$ for the Bretton Woods period, and in the post-Bretton Woods period it rose 47% to 3.95. Recalling that the variance of the real exchange rate rose 102% between the two periods, this means that the change in dynamics summarized in Ψ explains a bit under one half of the rise in the variance of the real exchange rate, with the remainder explained by the rise in exogenous innovations to the equation. While less than the share computed from the first-order autoregressive estimates, this still indicates an important role played by dynamics in generating greater real exchange rate volatility in the post-Bretton Woods period.

We apply the same method to derive the decomposition for monthly data. With 5 lags it is not possible to present a tractable analytical expression analogous to equation (4), but after the autoregressive parameter estimates are substituted, the numerical value of Ψ rises from 19.30 during

Bretton Woods to 52.37 afterwards, a rise of 271%. The variance of the actual real exchange rate data in our sample, measured at a monthly frequency, rose by 453% after the

[19] Our analysis is based on using real exchange rate data in level form rather than first differenced or filtered data. Given that we have already rejected the null hypothesis of nonstationarity of the real exchange rate for this sample, it is appropriate to use the levels data.

Bretton Woods period. Together this implies that 60% of the rise in variance of the real exchange rate data is due to a rise in the intrinsic dynamics, which is very similar to the results for annual data above.

Our results differ from Taylor (2002), who found only a modest increase in the half-lives of countries between the Bretton Woods period and the floating period that followed. He reports estimates pooling across countries as well separately for each country; since our panel methodology implicitly pools across countries, the pooled estimates in Taylor are the most appropriate comparison. Taylor's pooled estimates of the half-life rose modestly from 1.5 to 2.1 years. (When he estimated separately by country, the median half-life among his sample countries rose from 2.1 years to 2.6 years, while the mean rose only from 2.4 to 2.6 years.) Taylor graciously provided his dataset to us, enabling us to investigate the differences in results. The second row of Table 4 reports that when we apply our estimation methodology to Taylor's dataset, we find the same values as did Taylor: we estimate a half-life of 1.5 for Bretton Woods, and a half-life of 2.1 for the floating period.[20] We conclude that our particular methodology, Pesaran's CCEP panel estimation with bias correction, is not the source of difference in our main results.[21]

Next we trace where the difference does come from. Our dataset studies only developed economies, whereas Taylor includes three developing economies: Argentina, Brazil and Mexico. Each of these developing economies suffered a period of extremely high inflation, and such periods are known to be associated with unusually high speeds of adjustment in prices and exchange rates, which would tend to lower half-lives of the real exchange rate. Row 3 in Table 4

[20] For comparison, we follow Taylor (2002) in reporting values to one decimal place.
[21] This should also be reassuring to Taylor, as our methodology has the advantage of controlling for contemporaneous correlations of error terms in the cross-section, which can be a source of significant bias in country pair panel datasets.

shows that when we exclude these three developing countries from Taylor's sample and continue to apply our estimation methodology, we find a higher half-life for the floating period, 3.1 years, compared to that for Bretton Woods, 1.9 years.[22] This gap is widened yet further when we extend Taylor's sample period to 2011 as in our dataset, with half-lives of 1.9 for Bretton Woods and 3.5 for the floating period. These changes in the dataset are sufficient to explain a sizeable gap between the half-lives of the two periods.

As shown in the table, the remaining difference between these estimates and our benchmark results is due (1) to the fact that our dataset uses period averages to measure the exchange rate, as is common practice in the literature, rather than trying to find end-of-period values, and (2) the fact our data from IFS starts in 1949 rather than 1946 as in Taylor's data set.[23] The first of these points supports the conjecture in Taylor (2001) that the common practice of using end period values for the exchange rate, due to data limitations in the IFS dataset, can affect estimates of half-lives. Nonetheless, our estimates in Table 4 indicate that finding a large gap in half-lives between the two time periods does not depend upon this convention.

Our main result also differs from Sarno and Valente (2006), who find *faster* convergence under fixed exchange rate regimes. However, this difference appears to arise from the fact they report results combining the Bretton Woods period together with the Gold Standard, whereas we focus on Bretton Woods. While Sarno and Valente (2006) do not report half-lives separately for Bretton Woods separate from the Gold Standard, Taylor does report separate results for each regime. His estimates pooling across countries indicate a half-life twice as long under the Gold Standard as under Bretton Woods, 3.1 years versus 1.5 years. This suggests that the relatively

[22] We note that in Taylor's non-pooled estimates for each country separately, when one takes an average excluding the three developing economies, this raises the half-lives of both periods, but without creating a difference between periods. We conclude that the three developing economies must have a large effect on the pooled estimate in comparison to the simple average of the non-pooled estimates with each country receiving equal weight.
[23] Taylor's dataset for these extra years appear to be based upon interpolation of missing values.

higher half-life in Sarno and Valente could arise from the fact they chose to aggregate over these two fixed-exchange rate periods.

IV. Conditional on Shock: Vector Error Correction Estimation

We now investigate the source of the increase in persistence, using a vector error correction model. This approach permits us to decompose the real exchange rate into its two underlying components, the nominal exchange rate and the price level differential, as well as to identify the shocks to each component.[24] Mussa conjectured that the rise in short-run real exchange rate volatility under flexible exchange rates might have arisen from an increase in the prevalence of shocks to the nominal exchange rate permitted by such a policy regime. We conjecture this same phenomenon may have contributed to the rise in real exchange rate persistence. If the dynamics of real exchange rate adjustment differ depending on which of the two shocks is at work, and if the mix of these two shocks differs by policy regime, this could explain the change in the observed half-life of the overall real exchange rate.

The stationarity of real exchange rates implies cointegration of nominal exchange rates $(e_{j,t})$ and relative prices $(p_{j,t})$ with the cointegrating vector (1, 1). The adjustment process of nominal exchange rates and relative prices can be studied using the following panel vector error-correction model (VECM):

$$\Delta e_{j,t} = \alpha_{e,j} + \rho_{e,j}(q_{j,t-1}) + \mu^{e}_{e,j}(\Delta e_{j,t-1}) + \mu^{p}_{e,j}(\Delta p_{j,t-1}) + \zeta^{e}_{j,t} \tag{5a}$$

$$\Delta p_{j,t} = \alpha_{p,j} + \rho_{p,j}(q_{j,t-1}) + \mu^{e}_{p,j}(\Delta e_{j,t-1}) + \mu^{p}_{p,j}(\Delta p_{j,t-1}) + \zeta^{p}_{j,t}. \tag{5b}$$

This two-equation system decomposes the real exchange rate, $q_{j,t}$, into its two components, the nominal exchange rate, e_j, and the relative price level, p_j. It regresses the first difference of

[24] We employ this methodology in Bergin, Glick, and Wu (forthcoming), which documents the asymptotic properties of this estimator for an vector error correction model and describes a bootstrapped bias correction approach suggested by Kilian (1998). Our results employ this bias-corrected estimation methodology.

each of these components on the lag level of the real exchange rate, which summarizes the degree to which the data deviate from PPP. Other regressors in (5) control for level effects and short-run dynamics of the variables. The coefficients $\rho_{e,j}$ and $\rho_{p,j}$ reflect how strongly the exchange rate and prices respond to PPP deviations. Because negative movements in these variables work to reduce PPP deviations, they provide a measure of the speed of adjustment of nominal exchange rates and relative prices, respectively. To allow for possible cross section dependence in the errors, we computed CCEP estimators of the parameters by including as regressors the cross section averages of all variables ($\Delta\bar{e}_t$, $\bar{q}_{t-1}$, $\Delta\bar{e}_{t-1}$, and $\Delta\bar{p}_{t-1}$) and ($\Delta\bar{p}_t$, $\bar{q}_{t-1}$, $\Delta\bar{e}_{t-1}$, and $\Delta\bar{p}_{t-1}$) for the $\Delta e_{ij,t}$ and $\Delta p_{ij,t}$ equations, respectively.

Table 5 reports speed of adjustment coefficients on the lagged real exchange rate in each equation. Estimates are all significant at the 5% level, which supports the presence of co-integration between *e* and *p*, and the interpretation of equation (4) is a VECM. For both regime periods the response coefficient of the nominal exchange rate to PPP deviations, $\rho_{e,j}$ is much larger than the response coefficient of prices, $\rho_{p,j}$. For the post-Bretton Woods period the exchange rate response coefficient is 3 times larger than the price coefficient (-.168 vs. -.045); for the Bretton Woods period the factor is 2 (-.223 vs. -.102). This distinction between exchange rate and price responses lends support to the relevance of decomposing the real exchange rate into these two components.

The VECM also provides a basis for identifying distinct shocks to the system. We use a Cholesky ordering of the variables *e* then *p,* which identifies as an exchange rate shock any innovation in the nominal exchange rate that is not explained as an endogenous response to the lagged values in the regression equation (5a). A price shock is then identified as an innovation in the price level not associated with a contemporaneous movement in the exchange rate. This is the identification strategy used in Bergin, Glick, and Wu (forthcoming). It has several advantages

in the present context over the alternative ordering. First, it avoids imposing an assumption of price stickiness (implying no contemporaneous movement in price), but rather allows the data to speak about the degree of price rigidity in response to shocks. Second, as we will show in robustness checks below, the alternative ordering, where contemporaneous comovements of the nominal exchange rate and price shocks are classified as price shocks, leads to price shocks that appear to be contaminated by exchange rate shocks.

Figures 3 and 4 plot impulse responses to the two shocks in both sample periods, and Table 6 computes the half-lives of the real exchange rate for each shock based on these impulse responses. Recall from the single-equation autoregression results in Table 2 the finding that the half-life in the Bretton Woods period was roughly half that during the post-Bretton Woods period (2.27 vs. 4.31). As shown in Table 6, the difference is much smaller once we decompose by shock. Conditioning on an exchange rate shock, the half-lives of the real exchange rate for the two periods are 2.38 and 3.57 years, respectively. Also, conditional on a price shock, the half-lives for both periods are similar and both quite small, 0.61 and 0.54 years, respectively. Among these four numbers, the contrast between the two periods seems small in comparison to the contrast between the two shocks. In other words, the half-life in response to nominal exchange rate shocks is fairly long regardless of regime period, whereas that to price shocks is short regardless of regime.

The impulse responses in Figures 3 and 4 illustrate this point, as well as offering additional lessons about the mechanism of adjustment in the real exchange rate. Figure 3 shows that the real exchange rate adjusts gradually in response to a nominal exchange rate shock, both for the Bretton Woods and post Bretton Woods periods, whereas Figure 4 shows the real exchange rate adjusts very quickly to a price shock in both regimes. In addition, note in Figure 3 that the path of adjustment in the real exchange rate seems to mirror the adjustment in the nominal exchange rate, with very little movement in relative prices. This offers some support to the idea that prices are sticky in responding to nominal exchange rate shocks, particularly given

that our choice of identification assumptions did not impose any such restriction on the ability of price to respond contemporaneously. In the case of price shocks, there appears to be significantly more movement in the price level, with the price component contributing to the adjustment path of the real exchange rate.[25]

These lessons regarding the mechanism of adjustment can be formalized using the decomposition of Cheung et al (2004). Defining the impulse response of variable *m* to shock *n* as $\psi_{m,n}(t)$, note that $\psi_{q,n}(t) = \psi_{e,n}(t) + \psi_{p,n}(t)$. Then $g^q_{e,n}(t) = \Delta\psi_{e,n}(t) / \Delta\psi_{q,n}(t)$ measures the proportion of adjustment in PPP deviations explained by nominal exchange rate adjustment, and $g^q_{p,n}(t) = \Delta\psi_{p,n}(t) / \Delta\psi_{q,n}(t)$ measures the proportion explained by price adjustment, such that $g^q_{e,n}(t) + g^q_{p,n}(t) = 1$. The results in Table 7 indicate that adjustment of the real exchange rate to nominal exchange rate shocks at most horizons is accomplished through changes in the nominal exchange rate. An exception is that at a horizon of 10 years a greater fraction of adjustment occurs through price movement than through the exchange rate, but this number has little meaning, since figure 1 shows that the total amount of adjustment occurring at this horizon is very small in magnitude. Adjustment to price shocks is attributed nearly equally to adjustment in the nominal exchange rate and the price level. Again, the overall conclusion is that once we decompose by shock, the quantitatively important differences are attributable to the type of shocks, not the nature of the exchange rate regime.

All this brings us back to the question of how it can be that the overall half-life of the real exchange rate from the single equation autoregression could be so much larger in the post-Bretton Woods period than in the Bretton Woods period. One conjecture is that nominal exchange rate shocks were more prevalent in the later, flexible exchange rate period. Given that

[25] Appendix Figures A1-A2 report impulse responses for the alternative identification ordering (*p*, then *e*). For exchange rate shocks, the impulse response looks almost the same as for the benchmark identification. The impulse response to price shocks looks different. Oddly, in the post-Bretton Woods period the contemporaneous response of the exchange rate is actually larger than the movement in price, which calls into question the naming of such a shock as a price shock. This suggests the alternative identification may lead to contamination of price shocks with exchange rate shocks, probably because the latter are more frequent and larger than the former.

exchange rate shocks imply more persistent real exchange rate dynamics than do price shocks, the greater prevalence of exchange rate shocks in the post-Bretton Woods period would imply greater persistence in the real exchange rate overall.

Table 8 reports variance decompositions, which confirm this conjecture. The variance of real exchange rate forecast errors is almost entirely due to nominal exchange rate shocks during the post-Bretton Woods period . During the Bretton Woods period, exchange rate shocks account for 76% to 87% of variability in the real exchange rate.[26] While this is a surprisingly substantial share given that this is supposed to be a fixed exchange rate regime, it nonetheless is smaller in every year than the very large shares close to 100% for the flexible exchange rate regime period.

Our finding that exchange rate shocks explain a great proportion of real exchange variability during the post-Bretton Woods period is similar in spirit to the claim by Mussa and others that a flexible exchange rate regime allows more nominal exchange rate shocks, generating higher volatility. However, we note that our analysis generalizes and extends the argument of Mussa in two important respects. First, we focus on persistence as well as on the volatility of exchange rates, showing a new stylized fact regarding persistence. And second, we identify nominal exchange rate shocks separately from price shocks using formal econometric method s. This reveals that the primary driver of the contrasting behavior of real exchange rates across the two regimes is the contrasting dynamics of the two shocks in combination with the greater role of exchange rate shocks in the later period.

Lastly, we note that we intentionally do not take a stand on the possible sources for exchange rate shocks in our analysis. The rise in nominal exchange rate shocks may arise from money supply shocks unleashed by the relaxation of policy constraints during the Bretton Woods period. But they may also represent shocks to capital flows or to risk premia in the foreign exchange market. We leave it to future research to further decompose our nominal exchange rate

[26]Variance decompositions for the alternative identification ordering (*p* then *e*) are reported in the appendix. For the Bretton Woods period, the exchange rate and price shocks split almost exactly 50-50, with extremely wide confidence bands. This is further evidence of the inability to identify distinct shocks under the alternative identifying strategy.

shocks into these alternative sources, avoiding the inherently controversial identification strategies necessary to do so. Our main point regards the distinction between exchange rate and price shocks, as this distinction is found to imply quantitatively large and economically important differences in real exchange rate dynamics.

V. Conclusions

In this study we estimate how exchange rate regimes affect the long-run dynamics and half-life of the real exchange rate. We obtain two main results. First, we find that the speed of convergence of the real exchange rate to PPP is slower in the post-Bretton Woods period. Associated with the result, we find that up to two-thirds of the rise in measured short-run volatility during the more post-Bretton Woods period can be attributed to this increase in persistence. Second, when we identify nominal exchange rate and price shocks to the real exchange rate separately, we find that the half-lives and dynamics of the real exchange rate are essentially the same across regimes, and instead the key distinction is the differential response to these two shocks.

Our analysis offers a reinterpretation of the classic finding of Mussa (1986) concerning the increase in real exchange rate volatility after the demise of the Bretton Woods system. The overall conclusion is that the rise in real exchange rate volatility observed by Mussa worked substantially through a rise in persistence, which in turn was driven by the fact that nominal exchange rate shocks are more persistent than price shocks.

These findings have implications for the ongoing and fundamental debate over the source of real exchange rate fluctuations. Mussa used the fact that short-run real exchange rate volatility as affected by the nominal exchange rate regime to argue in favor of nominal shocks as an important driver of real exchange rates. In the same way we counter recent arguments against the role of nominal shocks in explaining real exchange rate persistence by pointing to the

observation that real exchange rate dynamics are also affected by the nominal exchange rate regime.

The findings also suggest an approach to resolving the "PPP puzzle:" the difficulty in explaining together the facts of high volatility and high persistence of real exchange rates. Monetary models with sticky prices are good at explaining real exchange rate volatility but not its persistence, whereas real models are good at explaining real exchange rate persistence but not volatility. However the finding above indicates that the two facts are best viewed as complementary: a high level of persistence helps create a high level of variance in the observed real exchange rate data. Our analysis suggests a role for dynamics in explaining both facts, and suggests that research focus on searching for a model that explains persistence as a means of explaining volatility.

Our findings also have lessons for the broader literature studying PPP convergence, suggesting that the literature should adopt a convention of conditioning results by shock. A standard univariate study of the real exchange rate will tend to conflate the distinct shocks to nominal exchange rates and price components of the real exchange rate, which seem to have very disparate dynamics.

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Table 1: Stationarity of the Real Exchange Rate

Sample	b	t-stat	# pairs	significance 1%	5%	10%
Annual data:						
Bretton Woods	-0.3322	-2.138	20	No	No	Yes
Post-Bretton Woods	-0.2397	-2.132	20	No	No	Yes
Whole sample	-0.1742	-2.496	20	Yes	Yes	Yes
Monthly data:						
Bretton Woods	-0.0363	-0.941	17	No	No	No
Post-Bretton Woods	-0.0193	-2.327	17	No	Yes	Yes
Whole sample	-0.0154	-2.303	17	No	Yes	Yes

Note: For annual data, the table reports estimates of the equation

$$\Delta q_{j,t} = a_j + b_j(q_{j,t-1}) + c_j(\bar{q}_{t-1}) + d_j(\Delta\bar{q}_t) + \varepsilon_{j,t} \quad ; \quad j = 1,...,N \text{ and } t = 1,...,T$$

where $\bar{q}_t = \sum_{j=1}^{N} q_{j,t}$ is the cross-section mean of $q_{j,t}$ across country pairs and $\Delta\bar{q}_t = \bar{q}_t - \bar{q}_{t-1}$.

For monthly data, the estimated equation is

$$\Delta q_{j,t} = a_j + \sum_{m=1}^{M} b_{j,m}(q_{j,t-m}) + \sum_{m=1}^{M} c_{j,m}(\Delta\bar{q}_{t-m}) + d_j\Delta\bar{q}_t + \varepsilon_{j,t} \quad ; \quad j = 1,...,N \text{ and } t = 1,...,T$$

where M =12.

Table 2. Half-lives in Autoregressions of Real Exchange Rates

Sample	#pairs	Half-life
Annual Data (one lag):		
Bretton Woods	20	2.27
(5%, 95%)		(1.43, 3.39)
(10%, 90%)		(1.57, 3.02)
Post-Bretton Woods	20	4.31
(5%, 95%)		(3.44, 5.24)
(10%, 90%)		(3.63, 4.97)
Annual Data (two lags):		
Bretton Woods	20	2.75
(5%, 95%)		(1.65, 4.71)
(10%, 90%)		(1.78, 3.98)
Post-Bretton Woods	20	4.27
(5%, 95%)		(3.51, 5.09)
(10%, 90%)		(3.68, 4.88)
Monthly Data (5 lags)		
Bretton Woods	17	2.23
(5%, 95%)		**(1.05, 4.49)**
(10%-90)		(1.19, 3.50)
Post-Bretton Woods	17	5.22
(5%, 95%)		(4.09, 6.50)
(10%, 90%)		(4.30, 6.19)

Note: Estimates are based on the equation $q_{j,t} = c_j + \sum_{m=1}^{M} \rho_{j,m}(q_{j,t-m}) + \varepsilon_{j,t}$ along with cross-sectional means of all dependent variables. Half-life in years are calculated from simulated impulse responses derived from the parameter estimates. Bias correction is carried out via the Kilian (1998) bootstrap method with 1000 iterations.

Table 3: Decomposition of Change in Real Exchange Rate Volatility

Sample	(1) var(*q*)	(2) ρ	(3) $\frac{1}{1-\rho^2}$	(4) (3)/(1)
Bretton Woods	0.012	0.737	2.188	--
Post-Bretton Woods	0.025	0.852	3.639	--
%Change	102.4%	--	66.4%	0.648

Note: Column (1) reports the average variance of the real exchange rate in the data set. Column (2) reports estimates of the autoregressive parameters from equation (2). Columns (3) and (4) are author computations based on the preceding columns.

Table 4: Robustness Checks Using Data from Taylor (2002)

	Half-life: Bretton Woods	Half-life: Floating Period
Taylor data and methodology (as reported in his paper)	1.5	2.1
Our Methodology:		
Taylor dataset	1.5	2.1
Excluding developing countries	1.9	3.1
Also extending data to 2011	1.9	3.5
Also using period averages	1.5	5.1
Also using our regime dates	2.8	4.7

Note: Half-lives are in years.

Table 5: Vector Error Correction Estimates

Sample	#obs.	Exchange Rate Equation	Relative Price Equation
Bretton Woods	20	-0.223	-0.102
(5%, 95%)		(-0.436, -0.111)	(-0.194, -0.030)
(10%, 90%)		(-0.397, -0.149)	(-0.174, -0.045)
Post-Bretton Woods	20	-0.168	-0.050
(5%, 95%)		(-0.221, -0.128)	(-0.069,-0.034)
(10%, 90%)		(-0.206, -0.136)	(-0.065, -0.038)

Note: The table reports estimates for the system:

$$\Delta e_{ij,t} = \alpha_{ij,e} + \rho_{e,ij}(q_{ij,t-1}) + \mu^{e}_{e,ij}(\Delta e_{ij,t-1}) + \mu^{p}_{e,ij}(\Delta p_{ij,t-1}) + \zeta^{e}_{ij,t}$$

$$\Delta p_{ij,t} = \alpha_{ij,p} + \rho_{p,ij}(q_{ij,t-1}) + \mu^{e}_{p,ij}(\Delta e_{ij,t-1}) + \mu^{p}_{p,ij}(\Delta p_{ij,t-1}) + \zeta^{p}_{ij,t}$$

along with cross-sectional means of all dependent variables. Bias correction is carried out via the Kilian (1998) bootstrap method with 1000 iterations. The t-statistics are computed from standard errors derived using the double-bootstrap method of Kilian (1998).

Table 6. Estimates of Half-lives of Real Exchange Rate Conditional on the Shock

	e shock	*p* shock
Bretton Woods	2.38	0.61
(5%,95%)	(0.81, 5.04)	(0.41, 0.92)
(10%, 90%)	(1.27, 3.88)	(0.44, 0.80)
Post--Bretton Wood	3.57	0.54
(5%,95%)	(2.80, 4.44)	(0.38, 0.70)
(10%, 90%)	(2.92, 4.24)	(0.40, 0.62)

Note: Half-lives in years, estimated from impulse responses of the equation system:

$$\Delta e_{j,t} = \alpha_{j,e} + \rho_{e,j}(q_{j,t-1}) + \mu_{e,j}(\Delta e_{j,t-1}) + \mu_{e,j}(\Delta p_{j,t-1}) + \zeta^{e}_{j,t}$$

$$\Delta p_{j,t} = \alpha_{j,p} + \rho_{p,j}(q_{j,t-1}) + \mu_{p,j}(\Delta e_{j,t-1}) + \mu_{p,j}(\Delta p_{j,t-1}) + \zeta^{p}_{j,t}$$

along with means of all dependent variables. The bias correction is carried out via the Kilian (1998) bootstrap method using 1000 replications.

Table 7: Relative Contributions of Nominal Exchange Rate and Price Adjustments to PPP Reversion

		exchange rate shock		price shock	
Bretton Woods:	years	$g^{q}_{e,e}$	$g^{q}_{p,e}$	$g^{q}_{e,p}$	$g^{q}_{p,p}$
	1	0.98	0.02	0.58	0.42
	2	0.73	0.27	--	--
	3	0.75	0.25	0.58	0.42
	5	0.65	0.35	0.47	0.53
	10	0.38	0.62	0.57	0.43
Post Bretton Woods:	years	$g^{q}_{e,e}$	$g^{q}_{p,e}$	$g^{q}_{e,p}$	$g^{q}_{p,p}$
	1	0.33	0.67	0.45	0.55
	2	0.93	0.07	0.25	0.75
	3	0.80	0.20	0.43	0.57
	5	0.82	0.18	0.47	0.53
	10	0.84	0.16	0.59	0.41

Note: The columns $g^{q}_{m,n}$ indicate the proportion of adjustment in the real exchange rate q explained by adjustment in variable m, conditional on shock n.

Table 8: Variance Decomposition of Real Exchange Rate

	years	exchange rate shock	price shock
Bretton Woods	1	0.76	0.24
	2	0.84	0.16
	3	0.86	0.14
	5	0.86	0.14
	10	0.87	0.13
Post-Bretton Woods	1	0.96	0.04
	2	0.98	0.02
	3	0.98	0.02
	5	0.98	0.02
	10	0.98	0.02

Note: Table reports the fraction of the forecast error variance of the real exchange rate due to each shock.

Fig. 1. Real Exchange Rate Impulse Responses, Autoregression

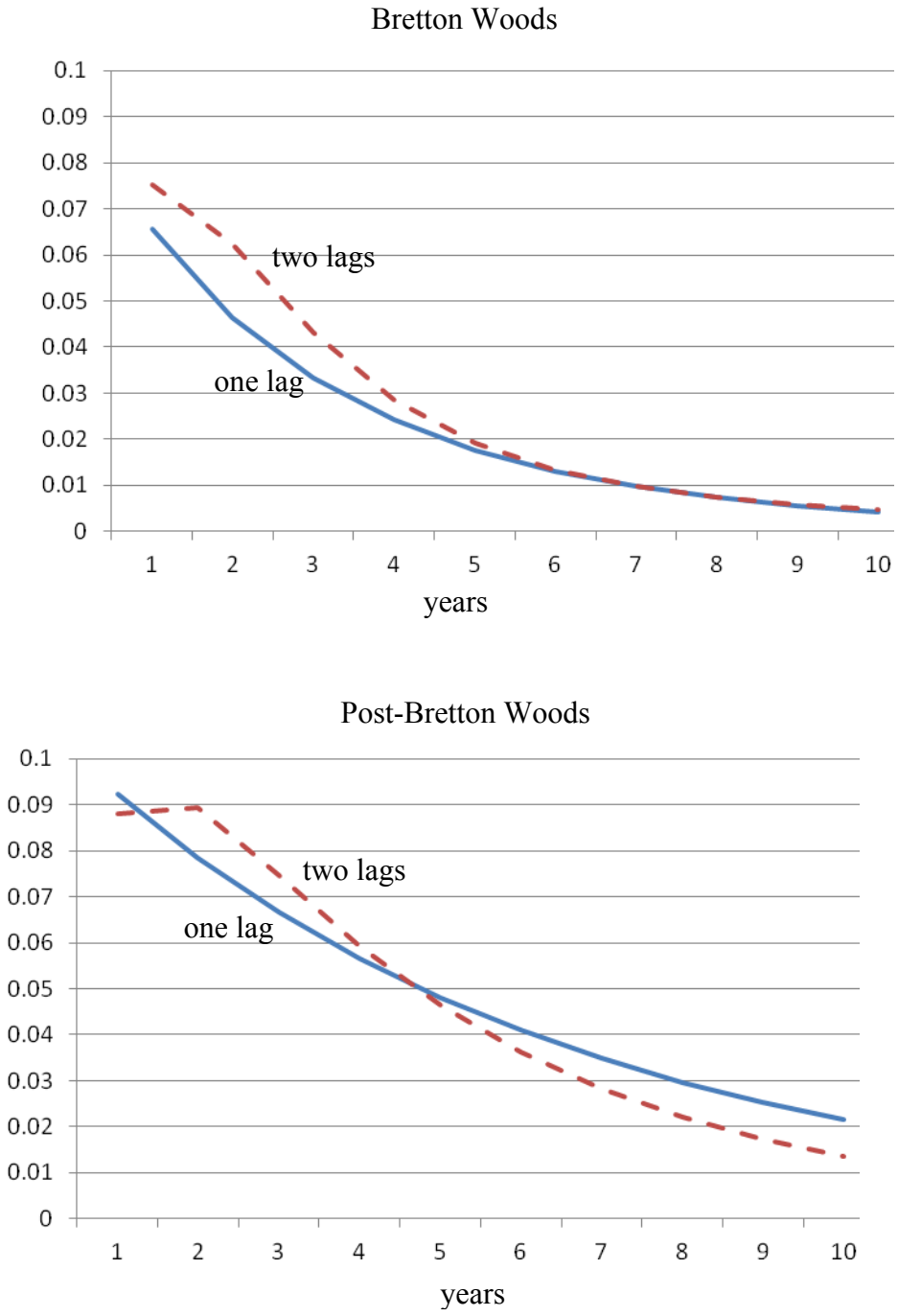
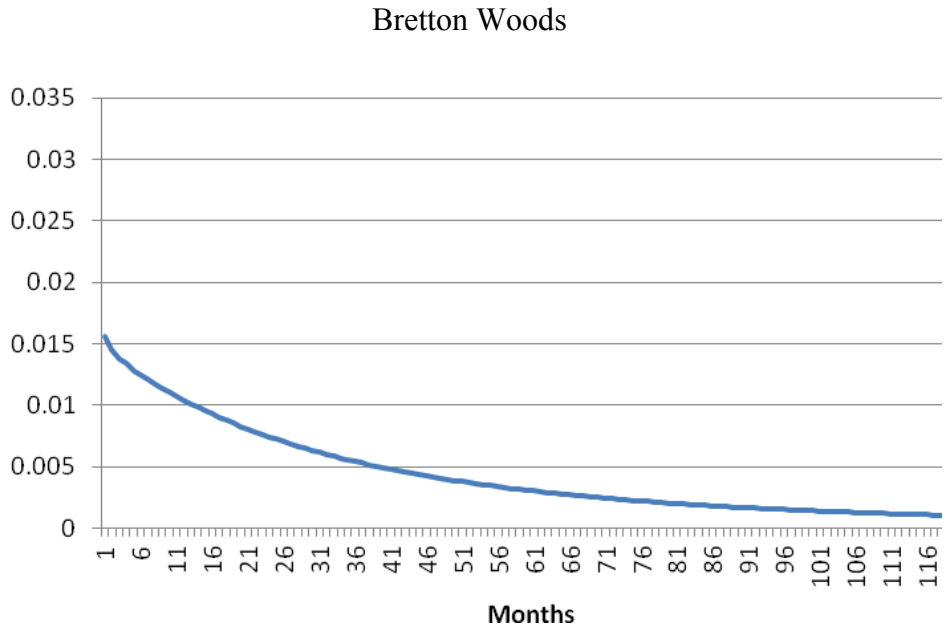


Fig. 1 plots impulse responses of the real exchange rate in annual data to a one-standard deviation shock to the real exchange rate, based upon CCEP estimates of the second-order autoregression version of equation (2) in the panel dataset. Vertical axes measure percent deviations; horizontal axis meaures years of the shock.

Fig. 2. Real Exchange Rate Impulse Responses,
Monthly Data, Autoregression with 5 lags



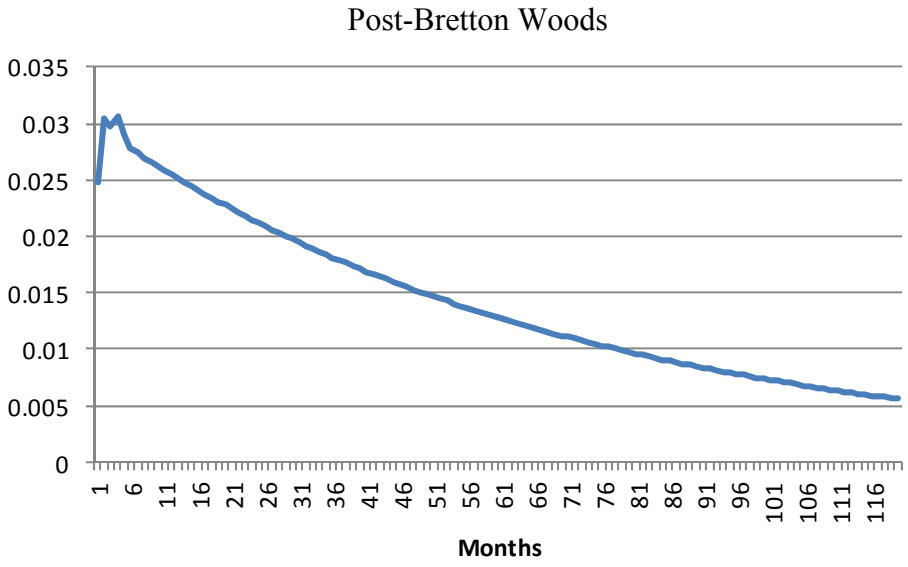
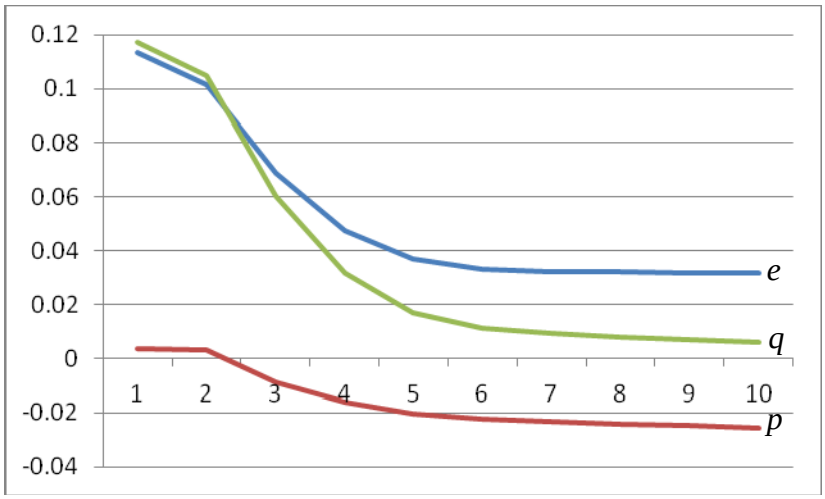


Fig. 2 plots impulse responses of the real exchange rate in monthly data to a one-standard deviation shock to the reral exchange rate, based upon CCEP estimates of the fifth-order autoregression version of equation (2) in the panel dataset. Vertical axes measure percent deviations.

Fig. 3. Impulse response to an *e* Shock

Bretton Woods



Post-Bretton Woods

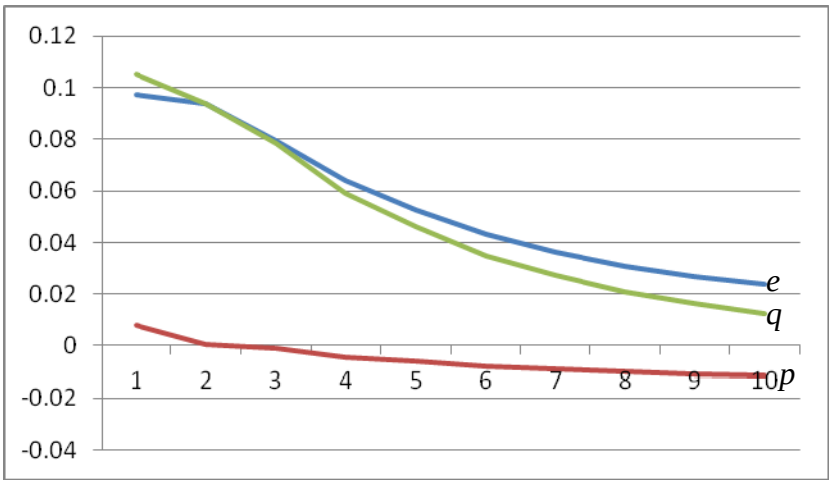
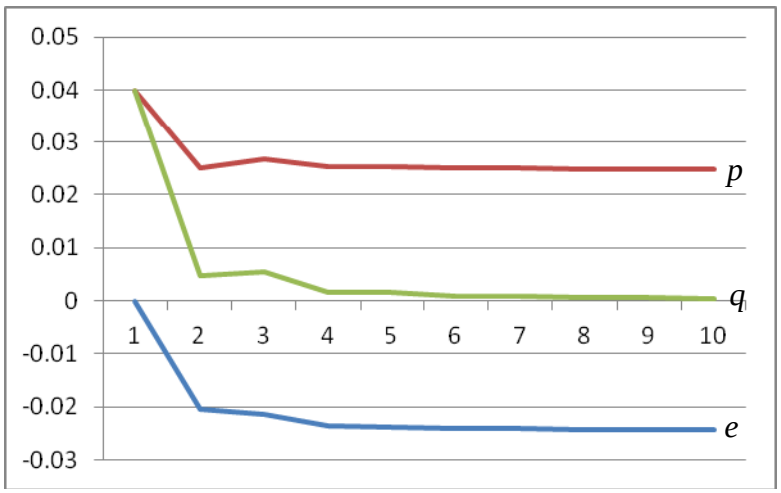


Fig. 3 plots impulse responses in annual data to a one-standard deviation shock to the nominal exchange rate, based upon CCEP estimates of the VECM model in equation (5) in the panel dataset. Vertical axes measure percent deviations. Horizontal axis measures years.

Fig. 4. Impulse Response to *p* Shock

Bretton Woods



Post-Bretton Woods

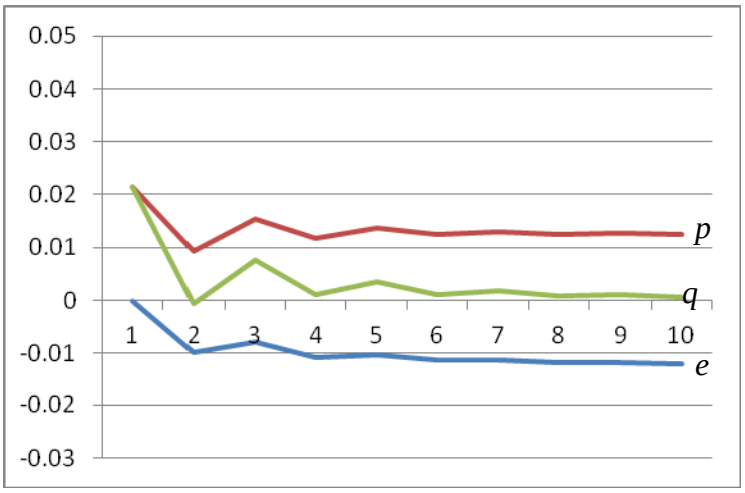


Fig. 4 plots impulse responses in annual data to a one-standard deviation shock to the ratio of national price levels, based upon CCEP estimates of the VECM model in equation (5) in the panel dataset. Vertical axes measures percent deviations. Horizontal axis measures years.

Appendix

Derivation of equation (4):

Rewrite the second-order autoregressive equation with autoregressive coefficients ρ_1 and ρ_2 as a two-equation first-order vector autoregression $y_t = Ay_{t-1} + \varsigma_t$, where $y_t = \begin{bmatrix} q_t & q_{t-1} \end{bmatrix}'$, $\varsigma_t = \begin{bmatrix} \varepsilon_t & 0 \end{bmatrix}'$, and $A = \begin{bmatrix} \rho_1 & \rho_2 \\ 1 & 0 \end{bmatrix}$. The covariance matrix of the residuals can be defined as $\Omega = E\left[\varsigma_t \varsigma_t'\right] = \begin{bmatrix} \text{var}(\varepsilon) & 0 \\ 0 & 0 \end{bmatrix}$. Following Chow (1980), the contemporaneous covariance can be written as follows:

$$E\left[y_t y_t'\right] = \Omega + \sum_{i=1}^{\infty}\left[BD^iB^{-1}\right]\Omega\left[BD^iB^{-1}\right]'$$

where *D* is the diagonal matrix of eigenvalues and *B* is the matrix of eigenvectors of *A*. This can be computed:

$$\Omega + BKB'$$

where the typical element (*i*,*j*) of *K* is

$$K_{ij} = \frac{\left(1-d_i\right)\left(1-d_j\right)M_{ij}}{1-d_id_j}$$

and where

$$M = B^{-1}\Omega[B^{-1}]'.$$

The variance of the real exchange rate is element (1,1) of the resulting matrix. This system was solved analytically for equation (4) using Matlab's symbolic toolbox.

Table A1. Estimates of Half-lives of Real Exchange Rate Conditional on the Shock, under Alternative Identification Ordering (*p* then *e*)

	e shock	*p* shock
Annual Data:		
Bretton Woods	1.93	1.83
(5%,95%)	(1.00, 3.36)	(0.62, 3.96)
(10%, 90%)	(1.18, 2.83)	(0.85, 2.88)
Post-Bretton Woods	3.11	3.43
(5%,95%)	(2.58, 3.71)	(2.56, 4.52)
(10%, 90%)	(2.69, 3.57)	(2.72, 4.17)

Note: Half-lives in years, estimated from impulse responses of the equation system:

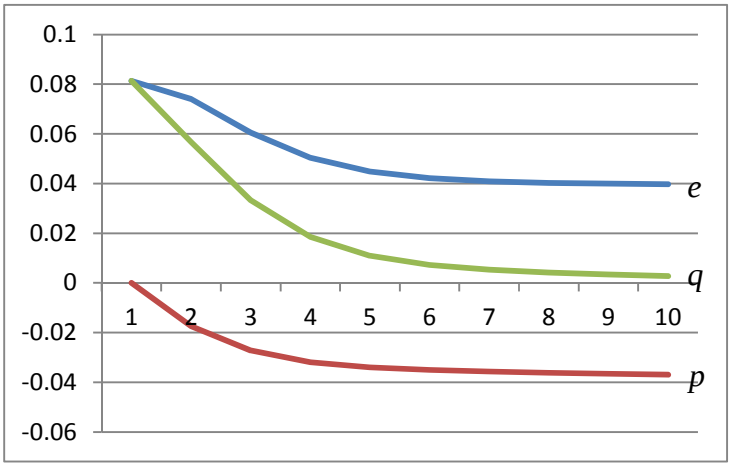
$$\Delta e_{ij,t} = \alpha_{ij,e} + \rho_{e,ij}(q_{ij,t-1}) + \mu^{e}_{e,ij}(\Delta e_{ij,t-1}) + \mu^{p}_{e,ij}(\Delta p_{ij,t-1}) + \zeta^{e}_{ij,t}$$

$$\Delta p_{ij,t} = \alpha_{ij,p} + \rho_{p,ij}(q_{ij,t-1}) + \mu^{e}_{p,ij}(\Delta e_{ij,t-1}) + \mu^{p}_{p,ij}(\Delta p_{ij,t-1}) + \zeta^{p}_{ij,t} .$$

The bias correction is carried out via the Kilian (1998) bootstrap method using 1000 replications.

Fig. A1. Impulse Response to an *e* Shock,
under Alternative Identification

Bretton Woods



Post-Bretton Woods

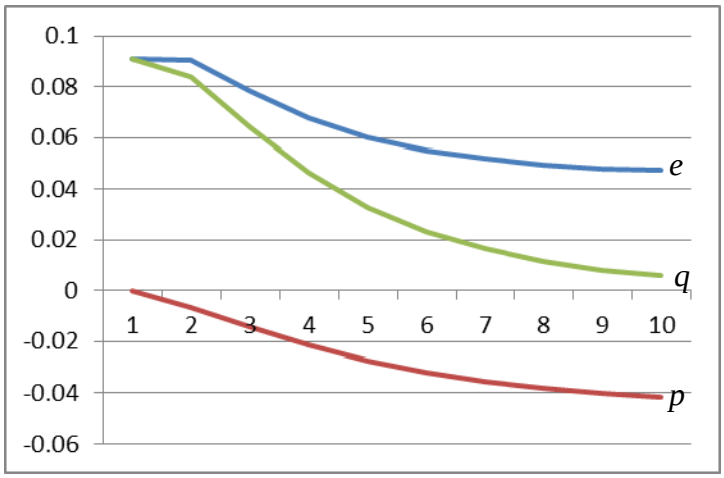
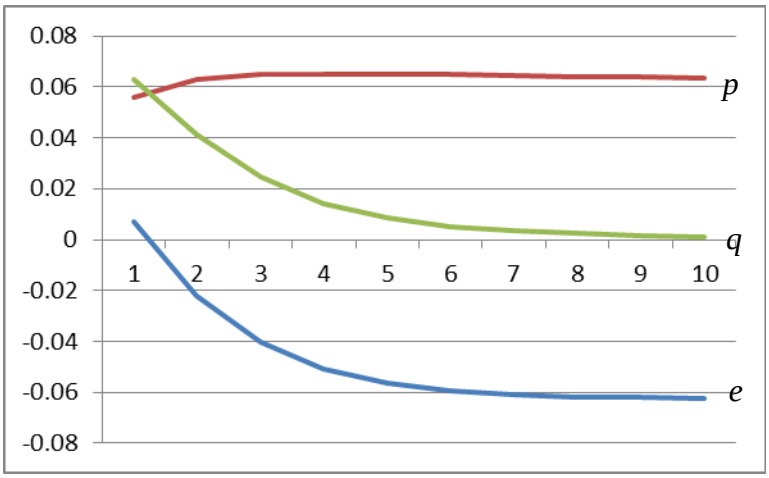
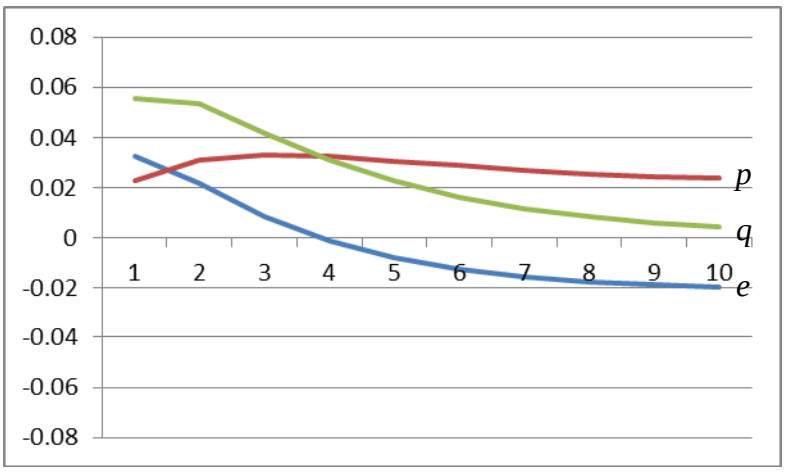


Fig. A2. Impulse Response to *p* Shock,
under Alternative Identification

Bretton Woods



Post-Bretton Woods