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ABSTRACT

While economic theory highlights the usefulness of flexible exchange rates in promoting adjustment in international relative prices, flexible exchange rates also can be a source of destabilizing shocks. We find that when countries joining the euro currency union abandoned their national exchange rates, the adjustment of real exchange rates toward purchasing power parity (PPP) became faster. To disentangle the possible causes for this finding we develop a novel methodology for conducting counterfactual simulations of an estimated VECM that distinguishes between the roles of the nominal exchange rate as an adjustment mechanism and as a source of shocks. We find evidence that prior to joining the euro currency union, member countries relied upon exchange rate adjustments as a mechanism to correct for PPP deviations arising from divergent domestic inflation rates. But the loss of the exchange rate as an adjustment mechanism after the introduction of the euro was more than compensated by the elimination of the exchange rate as a source of shocks, in combination with faster adjustment in national price indices. These findings support claims that flexible exchange rates are not necessary to promote long-run international relative price adjustment.

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1. Introduction

Economic theory has highlighted the ability of flexible exchange rates to promote adjustment in international relative prices towards equilibrium even when goods prices are sticky, a position famously championed in Friedman (1953). However, the foreign exchange market also can be a source of shocks, so that exchange rate flexibility may promote large and persistent deviations of the real exchange rate from long-run equilibrium. Indeed, as financial markets have become more integrated globally and international asset trade volume has grown larger compared to goods trade, nominal exchange rate fluctuations appear to be driven more by volatile financial market shocks than by pressure to balance relative goods prices.

The debate about the relative merits of exchange rate flexibility has played out prominently in arguments about the costs and benefits of joining the euro currency union, specifically whether the benefits of adopting a common currency exceed the costs of giving up the ability to promote equilibrium changes in the real exchange rate through nominal exchange rate adjustment. In contrast to Friedman's view, several recent papers have argued that the benefits of joining a currency union exceed the costs of sacrificing exchange rate flexibility. For example, Buiter (2008) argued that the "shock absorber" role of the exchange rate is quite limited and market-determined exchange rates are primarily a source of shocks and instability, implying that joining the euro would enable the United Kingdom to escape these destabilizing effects. More recently, Berka, Devereux, and Engel (2012) argued that the real exchange rate adjustment in a currency union like the eurozone might be superior to that under floating rates, both because exchange rates are disconnected from the foreign goods prices that consumers actually see, and because capital flows dominate nominal exchange rate movements.

This paper studies how adoption of the euro has affected the rate at which the real exchange rate of member countries adjusts to deviations from purchasing power parity (PPP).[2] We develop a novel methodology in order to distinguish between the roles of the nominal exchange rate, first, as a mechanism of adjustment to exogenous shocks, and, second, as a source of shocks. This methodology begins with estimating a vector error correction model (VECM) of the real exchange rate that decomposes the real exchange rate into the nominal exchange rate and the ratio of goods prices in local currency terms. This approach allows the exchange rate and prices to adjust at different speeds and also permits identification of shocks arising in the foreign exchange market separately from those in the goods market. We next estimate the half-life of the real exchange rate adjustment conditional on specific shocks, which we refer to as "conditional PPP."[3] We then conduct counterfactual simulations of the VECM system that mix and match individual parameters characterizing the pre-euro and euro periods, particularly parameters governing long-run and short-run dynamics. Comparing half-lives across these hypothetical scenarios allows us to measure the contribution of the exchange rate as a mechanism of adjustment separately from its contribution as a source of shocks.

We find that the rate at which the real exchange rate converges to its long run level became faster among European countries after they adopted the euro. This result is surprising, as we also find evidence that prior to the euro these countries indeed did rely upon nominal exchange rate adjustment to correct PPP deviations, including those deviations arising specifically from shocks to domestic goods prices. This empirical evidence is consistent with popular anecdotes of countries with higher than average inflation rates using currency devaluations to correct relative price

[2] There is a large and long-standing literature estimating rates of convergence to PPP. See Imbs, Mumtaz, Ravn, and Rey (2005) for a prominent example and discussion of this literature.
[3] Our VECM methodology shares elements with that in Bergin, Glick, and Wu (2014), but the analysis there did not seek to measure the relative contribution of the exchange rate as a source of shocks. It also focused on the effects of the Bretton Woods regime shift – not adoption of the euro -- on real exchange rate dynamics.

imbalances with European neighbors. Nevertheless, while the loss of this adjustment mechanism works to lengthen half-lives, we find it was more than compensated by two other factors working in the opposite direction. First, we find evidence that nominal exchange rate shocks were a substantial source of real exchange rate deviations among the countries in our sample prior to their adoption of the euro, and eliminating this source of persistent deviations under the euro lowered the average half-live of the real exchange rate. Second, we also find evidence that price adjustment in response to PPP deviations increased after the adoption of the euro. These two effects appear to have both worked to lower the half-live of the real exchange rate, and in combination they were more than enough to offset the loss of the exchange rate as an adjustment mechanism. In sum, we take these findings as support for claims that flexible exchange rates are not necessary to promote long-run international relative price adjustment.

In related literature, Cheung, Lai, and Bergman (2004) found that the speed of PPP convergence and real exchange rate persistence for several major currencies vis-a-vis the dollar during the floating rate period is driven largely by the behavior of the nominal exchange rate, with the exchange rate responding much more slowly than prices to shocks. However, in contrast to our analysis, they do not construct orthogonalized shocks to enable measurement of the relative contributions of exchange rate and price shocks. They also did not consider the effects monetary regime shifts, such as the adoption of the euro, on real exchange rate persistence.

Several papers have investigated PPP adjustment under the euro. Koedijk et al (2004), Lopez and Papell (2007), and Bahmani-Oskooee and Kutan (2008) conduct unit root tests of PPP, finding greater evidence of convergence for samples including the euro period. These papers, however, do not pursue explanations for this finding by estimating a VECM. In contrast to these other papers, Huang and Yang (2014) find that convergence is weaker after the introduction of the euro compared to earlier periods. While they do estimate a VECM, they do not condition by shock or use their VECM to run counterfactual simulations as we do to investigate the cause of the change in half-life.

In related work, Artis and Ehrmann (2006) compared the exchange rate as an adjustment mechanism and source of shocks using a different methodology, structural VARs. While their methodology offers a richer set of options for shock identification, it does not provide a formal metric of the contribution of exchange rate adjustment, as we do in terms of the half-life of the real exchange rate. They also do not employ panel techniques or conduct counterfactual simulations to distinguish alternative channels by which exchange rates matter.

The paper is organized as follows. The data and preliminary analysis involving stationarity tests of the real exchange rate for euro area members are presented in Section 2. The main empirical results are presented in the following two sections, with Section 3 estimating the half-live of the real exchange rate during the pre-euro and euro periods from single equation autoregressions, and Section 4 explaining the finding of a decline in half-life by estimating a VECM and conditioning on shocks. Section 5 presents conclusions.

2. Data

The dataset consists of consumer prices and bilateral nominal exchange rates for 9 original European member countries of the euro union with Germany as the numeraire, all taken from the *International Financial Statistics*.[4] The sample is monthly in frequency and covers the

[4] The full list of countries is: Austria, Belgium, Finland, France, Ireland, Italy, Netherlands, Portugal, and Spain. Luxembourg, also an original member of the euro union, is excluded. The price data are not seasonally adjusted.

period April 1973 to February 2014. The breakpoint between the pre-euro and euro periods is January 1999.

We define the real exchange rate, $q_{j,t}$ as the relative price level between country *j* and the base country (Germany) in period *t*, computed as $q_{j,t} = e_{j,t} + p_{j,t}$, where $e_{j,t}$ is the nominal exchange rate (German currency per currency *j*), and $p_{j,t} = p^*_{j,t} - p_{GER,t}$ is the log difference between the domestic price indices in country *j* and Germany and all variables are expressed in logs.[5] Hence, increases in *e* or *q* indicate nominal and real appreciation, respectively, of currency *j* against Germany's currency.

To check for stationarity, we apply the cross-sectionally augmented Dickey-Fuller (CADF) test suggested by Pesaran (2007) by estimating the panel regression:

$$\Delta q_{j,t} = a_{1j} + b_{1j}t + \omega_{0,j}q_{j,t-1} + \sum_{m=1}^{M-1}\omega_{1m,j}\Delta q_{j,t-m} + \omega_{2,j}\overline{q}_{t-1} + \sum_{m=0}^{M-1}\omega_{3m,j}\Delta\overline{q}_{t-m} + \varepsilon_{j,t}, \quad j = 1,...,N, \text{ and } t = 1,...,T \tag{1}$$

where $\overline{q}_t = \sum_{j=1}^{N} q_{j,t}$ is the cross-section mean of $q_{j,t}$ across the *N* country exchange rates, $\Delta\overline{q}_t = \overline{q}_t - \overline{q}_{t-1}$, and the purpose of augmenting the specification with cross-section means is to control for contemporaneous correlation among $\varepsilon_{j,t}$. The null hypothesis can be expressed as $H_0 : \omega_{0,j} = 0$ for all *j* against the alternative hypothesis $H_1 : \omega_{0,j} < 0$ for some *j*. The test statistic provided by Pesaran (2007) is given by:

$$CIPS(N,T) = (1/N)\sum_{j=1}^{N} t\text{-}stat_j(N,T),$$

[5] This specification assumes that $p_{GER,t}$, $p^*_{j,t}$ share similar convergence speeds, a property that has been found to be consistent with the data; see Cheung et al. (2004).

where $t\text{-}stat_j(N,T)$ is the t statistic of ω_{0j} from the estimation of equation (1). The AIC criterion is applied to select the appropriate lag order in equation (1).[6] Setting the maximum lag length of M to 12, the selected optimal lag length is 10 for the pre-euro period ($M = M_1$ =11) and 9 for the euro period ($M = M_2$ =10).[7]

Based on the selected lagged order, the unit-root hypothesis for q is rejected at the 5% level (*t-stat* = -2.883) for the pre-euro period. However, the unit-root hypothesis is not rejected at the 10% level (*t-stat* =-2.378) for the euro period. Pesaran (2007) indicates that the power of the CIPS test is low when the sample size (N and T) is small. Thus, a likely reason for failing to reject the unit-root hypothesis for the euro period is its small sample size.[8]

3. Estimating rates of convergence

To estimate the rate of real exchange rate convergence, we first estimate an autoregressive panel model that nests the data for the pre-euro and euro periods together:

$$q_{j,t} = d_{pre-euro,t}(a_{1j} + b_{1j}t + \sum_{m=1}^{M_1}\eta_{1m}q_{j,t-m}) + d_{euro,t}(a_{2j} + b_{2j}t + \sum_{m=1}^{M_2}\eta_{2m}q_{j,t-m}) + \varepsilon_{j,t}\,. \quad (2)$$

where the indicator regime variable $d_{pre-euro,t}$ takes a value of 1 during the pre-euro period and a value of 0 otherwise, i.e., $d_{pre-euro,t}$= 1 for t = 1,…, T_1, the end date of the pre-euro period, and

[6] More specifically, we impose the restriction that each country has the same lag order M and estimate equation (1) by OLS for sequentially longer lag lengths, i.e. M = 2, 3, … $M^{\max}$, where $M^{\max}$, the maximum lag length, is set to 12 in lieu of the monthly frequency of the dataset. The AIC value for each lag order is constructed based on the residual matrix that includes the estimated residuals for all countries in the panel. The optimal lag order is the specific lag order that results in the minimum AIC value.

[7] We obtain similar results for lag order from applying the AIC criterion to each country in the panel individually: the optimal lag lengths for Austria, Belgium, Finland, France, Ireland, Italy, Netherlands, Portugal, and Spain for the pre-euro period are 12, 7, 11, 7, 11, 2, 10, 11, 11 months, respectively, with a median of 11, and for the euro period are 11, 9, 7, 10, 12, 8, 9, 7, 10 months, with a median of 9.

[8] We proceed with the analysis despite the lack of a statistically significant rejection of the unit root hypothesis for the euro period. This choice appears to be vindicated in the analysis below, where we find in the full sample nesting both periods a statistically significant result indicating a faster rate of convergence for the real exchange rate during the euro period relative to the pre-euro period.

correspondingly $d_{euro,t} = 1$ for $t = T_1 + M_2 + 1, \ldots, T$ and 0 otherwise.[9] To control for contemporaneous correlation of residuals, the common correlated effects pooled (CCEP) regressor of Pesaran (2006) is used, involving augmentation of equation (2) with the cross-sectional means of dependent and explanatory variables during the two regimes.[10] To control for potential bias in the CCEP estimator from the presence of lagged dependent variables, the standard double bootstrap procedure of Kilian (1998) is employed with 1000 replications to obtain bias-adjusted estimates for each sub-period.[11] The optimal lag length in equation (2) is M_1=11 for the pre-euro period and M_2=10 for the euro period.

Table 1 reports coefficient estimates and half-lives of the real exchange rate, computed on the basis of simulated impulse responses.[12] The half-life estimated for the pre-euro period is 2.39 years (with a 5%-95% band of 1.81 to 3.69 years); that for the euro period is 1.35 years (with a band of 1.03 to 2.01 years). This represents a 44% drop in persistence in the euro period.[13] As a test of significance, our stochastic simulations also compute the difference in half-life (1.05) between the two regimes; these results are reported in the last column. The 5%-95%

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[9] The estimation start dates are adjusted for the number of lags, so that the data for lagged and contemporaneous variables are drawn consistently from the same subsample (euro or pre-euro periods). Thus the estimation period for the euro period begins at time $t = T_1 + 1 + M_2$.

[10] As discussed in Pesaran (2006), the cross-sectional means are observable proxies for the common effects in the panel that enter the $\overline{H}_w$ matrix in his formula for the CCEP estimator. STATA code to conduct CCEP estimations used throughout the paper are available upon request.

[11] See the appendix of Bergin, Glick, and Wu (2013) for a Monte-Carlo study of the bias of the CCEP estimator when applied to models with a lagged dependent variable. In implementing the Kilian (1998) procedure to control for potential estimator bias, we resample residuals (filtering out the constant and trend by currency pair for each regime period) with replacement, initialize with demeaned data, and discard the first fifty simulated observations to eliminate the initial value effect.

[12] 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 $(f(t_1) - 0.5))/(f(t_1) - f(t_1+1))$.

[13] The 5% and 95% confidence bands for the half-life are constructed from its bootstrap distribution.To construct this distribution, we first bootstrap estimated residuals and use them to generate a pseudo data series of real exchange rates. We then re-estimate (2) with CCEP using this pseudo data and compute the half-life accordingly. The bootstrap distribution is constructed with 2000 iterations, and we report the 5th and 95th percentiles of the half-lives from the constructed bootstrap distribution.

confidence band for this difference of 0.13 to 2.30 excludes zero. These estimates support the conclusion that the half-life of the real exchange rate is lower in the euro period.

Figure 1 plots the impulse response functions (IRF) of the real exchange rate together with 5% and 95% confidence intervals during the pre-euro and euro periods, respectively. Except for some fluctuations in the initial periods, the IRF decreases monotonically in both periods, with the rate of decline of the IRF greater in the euro period. This is consistent with the result of a shorter half-life of the real exchange rate since the adoption of the euro.[14]

The shorter half-life during the euro period is surprising, as theories dating back to Friedman (1953) posit that a flexible exchange rate should be useful as an adjustment mechanism for relative prices when nominal prices are relatively rigid. This suggests that the eliminating adjustment of the nominal exchange rate by joining a currency union should raise the half-life of the real exchange rate rather than lower it.

For the sake of completeness and for later reference, we also estimate the autoregression separately for each sample period, not nesting across periods. More specifically, we estimate:

$$\text{Pre-euro: } q_{j,t} = a_{1j} + b_{1j}t + \sum_{m=1}^{M_1} \eta_{1m}(q_{j,t-m}) + \varepsilon_{j,t}, \quad j=1,\ldots,N, \; t=1,\ldots,T_1, \tag{3.1}$$

$$\text{Euro: } p_{j,t} = a_{2j} + b_{2j}t + \sum_{m=1}^{M_2} \eta_{2m}(p_{j,t-m}) + \varepsilon_{j,t}, \quad j=1,\ldots,N, \; t=T_1+1+M_2,\ldots,T, \tag{3.2}$$

where $M_1 = 11$ for the pre-euro period, and $M_2 = 10$ for the euro period. Note that since the nominal exchange rate is effectively fixed during the euro period, $\Delta e = 0$, $\Delta q = \Delta p$, and $q = p$, when the log of the exchange rate e is normalized at 0, implying that estimation of the autoregression of q is equivalent to estimating the autoregression equation (3.2) in p during this

[14] The IRF appears excessively jagged during the euro period because the seasonality in relative prices shows up in real exchange rate changes when the exchange rate is fixed. The IRF during the pre-euro period is less jagged since the exchange rate is able to respond and partially offset the seasonal variation in relative prices.

period. We estimate the above equations along with cross-sectional means of the left-hand and right-hand variables ($\bar{q}_t$, $\bar{q}_{t-1}$,..., $\bar{q}_{t-M}$) for (3.1) and ($\bar{p}_t$, $\bar{p}_{t-1}$,..., $\bar{p}_{t-M}$) for (3.2). Compared to the specification in equation (2), equations (3.1) and (3.2) have the disadvantage of not allowing for direct tests of statistical significance for the change between periods, as well as some loss of efficiency due to the smaller sample sizes. However, in addition to providing a complement to our nested regression, the estimated coefficients from the non-nested regressions, particularly equation (3.2), are useful in simulation exercises described in section 4.

Table 2 reports parameters of the AR(11) estimated for the pre-euro period in the first column. The estimated half -life of the real exchange rate is 2.39 years, which is very close to the value estimated from the autoregression nesting both periods together. The second column of the table reports estimates of the AR(10) of p for the euro period, and an estimated half-life of 1.61 years, which is somewhat higher than that estimated from the AR nesting both periods together. Figure 2 plots the estimated IRF of the real exchange rate during the pre- and euro periods, respectively; the dynamics generally appear very similar to those in Figure 1. However, it should be noted that the magnitude of one-standard deviation shocks, and hence the initial impact of these shocks, differ between the nested and non-nested cases. The nested case reported in Figure 1 implicitly imposes the same standard deviation for q shocks (0.0112), during the pre-euro and euro periods. In contrast, the non-nested case considered in Figure 2 allows them to differ across periods and yields a one-standard deviation p (=q) shock (0.004) in the euro period that is much smaller (the standard deviation of q in the pre-euro period is 0.0137).

4. Decomposing the role of shocks and dynamics

We now investigate the source of the change in real exchange rate persistence, using a vector error correction model (VECM). This permits us to decompose the dynamics of the real exchange rate into that of its two underlying components, the nominal exchange rate and the relative national price levels.[15]

4.1. Estimation of a vector error correction model

The adjustment process of nominal exchange rates and relative prices during the pre-euro period can be studied using the following panel VECM:

$$\begin{bmatrix} \Delta e_{j,t} \\ \Delta p_{j,t} \end{bmatrix} = \begin{bmatrix} g_{10,j} \\ g_{20,j} \end{bmatrix} + \begin{bmatrix} g_{11,j} \\ g_{21,j} \end{bmatrix} t + \begin{bmatrix} \rho_{10} \\ \rho_{20} \end{bmatrix} q_{j,t-1} + \begin{bmatrix} a_{11} & b_{11} \\ c_{11} & d_{11} \end{bmatrix} \begin{bmatrix} \Delta e_{j,t-1} \\ \Delta p_{j,t-1} \end{bmatrix} + ... + \begin{bmatrix} a_{1M-1} & b_{1M-1} \\ c_{1M-1} & d_{1M-1} \end{bmatrix} \begin{bmatrix} \Delta e_{j,t-M+1} \\ \Delta p_{j,t-M+1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{ej,t} \\ \varepsilon_{pj,t} \end{bmatrix}. \quad (4)$$

This two-equation system decomposes the real exchange rate, $q_{j,t}$, into the nominal exchange rate, $e_{j,t}$, and the relative price level, $p_{j,t}$, and regresses the first difference of each of these components on the lagged level of the real exchange rate. The coefficients ρ_{10} and ρ_{20} reflect how strongly the exchange rate and prices each respond to PPP deviations. To the extent these coefficients are negative, they provide a measure of the speed of adjustment of nominal exchange rates and relative prices, respectively, in reducing PPP deviations. The other regressors in the VECM control for level effects and short-run dynamics of the variables. As with our previous autoregressions, in order to handle possible cross-section dependence in the errors, we compute CCEP estimators of the parameters in each equation by including as regressors the cross-section averages of the dependent variable, $q_{j,t-1}$ and the lags of $\Delta e_{j,t}$ and $\Delta p_{j,t}$. An optimal lag length of M = 10 is determined from the median of the optimal lag lengths

[15] We employ this methodology in Bergin, Glick, and Wu (2013), 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. As with the autoregression estimation, this involves bootstrapping by resampling of residuals (filtering out the constant and trend by currency pair) with replacement, initializing with demeaned data, and discarding the first 50 simulated observations to eliminate the initial value effect.

of vector autoregressions of $e_{j,t}$ and $p_{j,t}$ for individual countries over the pre-euro period.[16]

The VECM system can be estimated only for the pre-euro sample period, as there is (obviously) no nominal exchange rate adjustment during the euro period. Since only relative prices can adjust during this period, the specification reduces to the following AR equation for the relative price during the euro period:

$$\Delta p_{j,t} = a_{2j} + b_{2j}t + \gamma_0 p_{j,t-1} + \gamma_1 \Delta p_{j,t-1} + \gamma_2 \Delta p_{j,t-2} + ... + \gamma_{M_2-1} \Delta p_{j,t-M_2+1} + \varepsilon_{pj,t} . \quad (5)$$

The coefficients in (5) can be obtained by a simple transformation of the coefficients (η_{2m}) in equation (3.2), since $\gamma_0 = \sum_{m=1}^{M_2} \eta_{2m} - 1$ and $\gamma_m = -\sum_{j=m+1}^{M_2} \eta_{2j}$, for m=1, 2, …, $M_2 - 1$. Hence, the coefficient estimates for *p* reported for the euro period in Table 2 can be used to recover estimates of γ_m. Thus, for example, $\hat{\gamma}_0 = -0.040$.

Under appropriate parameter restrictions, the VECM can be seen to nest equation (5). We will use the VECM estimated over the pre-euro period to measure the effect on the half-life of the real exchange rate of various counterfactual exercises, including removing the nominal exchange rate as a source of shocks, and removing the nominal exchange rate as an adjustment mechanism, in order to gauge how much these two factors contributed to the change in the real exchange rate half-life found in the preceding section.

Table 3 reports VECM estimates and the half-life of the real exchange rate conditional on a one standard- deviation shock to the nominal exchange rate, the relative price level, and to both the nominal exchange rate and price level together, all during the pre-euro period. The results indicate that the estimated error-correcting coefficients ($\hat{\rho}_{10} = -0.022$, $\hat{\rho}_{20} = -0.01$) are both

[16] The individual VAR lags for Austria, Belgium, Finland, France, Ireland, Italy, Netherlands, Portugal, and Spain are 11, 2, 12, 2, 10, 12, 10,4, 2, 10 months, respectively, with a median of 10.

negatively significant, indicating that the exchange rate and prices respond to PPP deviations. Examination of the short-run dynamics indicates that $\Delta e_{j,t}$ and $\Delta p_{j,t}$ each depend more on their own lags.

The VECM provides a basis for identifying distinct shocks to the system. We use a Cholesky ordering of the variables *e,* then *p,* which identifies an exchange rate shock as any innovation in the nominal exchange rate that is not explained as an endogenous response to the lagged values in the exchange rate regression. A price shock is then identified as an innovation in the price level not associated with a contemporaneous movement in the exchange rate. This identification has an advantage in the present context in that it avoids imposing an assumption of price stickiness (implying no contemporaneous movement in prices), but rather allows the data to speak about the degree of price rigidity in response to shocks. This identification allows us to distinguish between how well PPP holds conditionally for different types of shocks, which we refer to as conditional PPP. We find that this distinction permits a deeper understanding of real exchange rate behavior in our dataset, as explained below.

Table 3 also reports the half-life of the real exchange rate conditional on specific shocks and on both shocks simultaneously. The conditional half-life is 2.33 years for an exchange rate shock, 1.14 years for a price shock, and 2.00 years when both shocks occur simultaneously. Thus, the half-life conditional on an exchange rate shock is larger than that for a price shock, and the conditional half-life of real exchange rates when both shocks occur simultaneously is closer to that for a nominal exchange rate shock.

Figure 3 plots impulse responses (IRFs) of the real exchange rate to exchange rate and price shocks in the pre-euro period. The IRF of the real exchange rate first increases and subsequently declines regardless of shock. The decrease of the IRF of *q* for a price shock is faster than that for an exchange rate shock. When both shocks occur simultaneously, the IRFs of *q* are very similar to that for an exchange rate shock. Hence, Figure 3 supports the characterization of the results reported in Table 3.

4.2. Counterfactual simulations of the VECM system

Our finding that the half-life of the real exchange rate in the euro period is significantly lower than that in the pre-euro period challenges the argument made by Friedman (1953) that eliminating adjustment of the nominal exchange rate in response to relative price differences should raise the persistence of real exchange rate shocks. In this section we investigate what factors may explain our finding with counterfactual simulations. We do so by using our VECM system (4) to examine how different calibrations of the dynamic coefficients affect the impulse response functions and the corresponding half-life of the real exchange rate.

These simulation cases are presented in Table 4. As a benchmark, simulation 1 reports half-lives conditional on nominal exchange rate shocks, on price shocks, and on draws taken simultaneously from both shocks, with all dynamic parameters set at their values estimated from the pre-euro period as given in Table 3. We report the last case with simultaneous exchange rate and price shocks to use for comparison with the AR estimates of the unconditional half-life of *q* reported for the non-nested regressions in Table 2. The results for simulation 1 in column 3 of Table 4 indicate that the VECM estimates imply a half-life of 2.00 years in response to simultaneous *e* and *p* shocks. This is somewhat lower than the unconditional half-life of 2.39 estimated for the pre-euro period using the *q* autoregression (as reported in Table 2). A difference between the VECM and AR estimates is not surprising, as the VECM allows for more free parameters than the AR, in particular by allowing parameters in the *e* and *p* equation to differ from each other.[17]

We next verify that the VECM can replicate the half-life of the real exchange rate during the euro period with an appropriate set of parameter restrictions. In simulation 2, we set all of the coefficients in the Δe equation to zero, and set all of the coefficients in the Δp equation to their values from equation (5) estimated during the euro period. The simulation makes use of the AR(10) estimates of equation (3.2) for *p* reported in Table 2 for the euro period to recover estimates of the γ_i coefficients in equation (5), as discussed in Section 4.1. Specifically, $\hat{\gamma}_0 = \sum_{m=1}^{M_2} \hat{\eta}_{2m} - 1$ and $\hat{\gamma}_m = -\sum_{j=m+1}^{M_2} \hat{\eta}_{2j}$, for *m*=1, 2, …, $M_2 - 1$. Simulation 2 generates a half-life conditional on *p* shocks of 1.60 years (see column 2 of Table 4), very close to the value of 1.61 estimated for *p* shocks during the euro period reported in Table 2. This result indicates that we can indeed capture the fall in half-life due to the introduction of the euro in terms of a specific set of parameter restrictions and identification of shocks. We proceed by assessing the relative contribution of each of these restrictions by imposing them individually in our simulations and conditioning on specific shocks.

We first consider the role that adoption of the euro played in eliminating the exchange rate as a source of shocks. Simulation 1 of Table 4 yields insight from the effect of restricting the source of shocks. Specifically, if exchange rate shocks are eliminated and only price shocks drive the real exchange rate, the half-life of *q* falls from 2.00 to 1.14, a 43% drop in the half-life. This magnitude decline is roughly the same percentage (44%) by which we found in Table 1 that

[17] Under the appropriate parameter restrictions, the VECM can of course replicate the real exchange rate dynamics of the AR equation estimated for *q* during pre-euro period. In particular, we estimate the following AR(10) equation for *q* with pre-euro data: $\Delta q_t = a_1 + b_1 t + \psi_0 q_{t-1} + \psi_1 \Delta q_{t-2} + .. + \psi_9 \Delta q_{t-9} + \varepsilon_{qt}$. We then impose the following parameter restrictions on the VECM coefficients: $\rho_{10} = \rho_{20} = 0.5\hat{\psi}_0$; $a_{1i} = b_{1i} = c_{1i} = d_{1i} = 0.5\hat{\psi}_i$, for $i = 0,...,9$. Simulation of the restricted VECM system implies an unconditional half-life for the real exchange rate of 2.36, which is the same as that derived from the non-nested estimate of an AR(10) of *q* for the pre-euro period. The above results are not reported in the paper but are available upon request.

the half-life fell during the euro period compared to the pre-euro period. Thus, this experiment suggests that the absence of exchange rate shocks during the euro period alone may potentially explain the fall in half-life of the real exchange rate.

Why is the effect of eliminating exchange rate shocks so powerful? Simulation 1 also indicates that the half-life of the real exchange rate is 2.33 years conditional on an *e* shock and only 1.14 years conditional on a *p* shock. Thus, there is a noticeable difference in the dynamics of the real exchange rate generated by the two shocks, with greater persistence associated with fluctuations arising from an *e* shock.[18] This reflects the more gradual *q* response and more persistent deviations in the real exchange rate observed for *e* shocks compared to *p* shocks shown in Figure 3.

The reason that exchange rate shocks during the pre-euro period lead to more persistent deviations in the real exchange rate lies largely with the fact that nominal exchange rates tend to exhibit significant delayed overshooting, that is, exchange rate changes grow for a period of time before diminishing. To show this, Figure 4 plots the IRF of the nominal exchange rate, real exchange rate, and relative prices, conditional on an *e* shock and a *p* shock, respectively. The delayed overshooting of the nominal exchange rate to an *e* shock (the dashed line) is indeed more significant than that to a *p* shock, in that the adjustment is longer. It takes about one and a half years for the nominal exchange rate to return to the level of the initial impact effect in the case of an *e* shock, but less than half a year in the case of a *p* shock.

Further insight can be gleaned by simulating a version of the VECM where the coefficients of short-run dynamics governing the response of exchange rate changes to lagged exchange rate changes are restricted to zero: $a_{11} = a_{12} = ... = a_{1,M-1} = 0$. In this case, as reported in

[18] Cheung et al (2004) also found greater persistence associated with the nominal exchange rate. However, they did not find any distinction in the half-life conditional on shock; instead they found that *q* adjustment due to nominal exchange rate adjustment was slower than adjustment due to the price component of the real exchange rate.

simulation 3 of Table 4, the half-life conditional on exchange rate shocks then falls to a level (1.71), nearly the same as that conditional on price shocks (1.64). Thus the short-run overshooting dynamics of the nominal exchange rate are what make *e* shocks lead to more real exchange rate persistence than *p* shocks.

While simulation 1 in Table 4 indicates that the absence of exchange rate shocks might be a significant factor explaining the lower half-life of *q* during the euro period, what about the loss of the "shock absorber" role of the exchange rate as an adjustment mechanism? In simulation 4 we run an experiment estimating the half-life in a world where the exchange rate is eliminated as a mechanism of adjustment, but remains a source of shocks. More specifically, simulation 4 imposes the restriction that $\rho_{10} = 0$, so that the nominal exchange rate does not respond directly to eliminate PPP deviations. In this case the estimated half-life balloons by a factor of four regardless of the shock on which one conditions (from 1.14 to 5.39 years for price shocks and 2.33 to 10.65 years for exchange rate shocks).

We draw several lessons from simulation 4. First, it provides evidence that in the pre-euro period European countries indeed did rely upon nominal exchange rate adjustments to correct for PPP deviations. Second, the fact that this is true regardless of the source of shocks suggests this adjustment was not simply a matter of the nominal exchange rate correcting itself after nominal exchange rate shocks. It appears that European countries relied upon exchange rate adjustment in response to shocks to goods prices as well. This is consistent with popular anecdotes of countries with higher than average inflation rates using currency devaluations to correct relative price imbalances with European neighbors. Third and most importantly, this effect, by implying greater persistence of the real exchange rate both in response to price as well as nominal exchange rate shocks, works in the opposite direction of explaining our primary finding that the introduction of the euro decreased real exchange rate persistence. Moreover, the finding that the ballooning of the half-life of *q* in this case occurs even when conditioning solely on price shocks indicates that the elimination of exchange rate shocks alone is insufficient to explain the decline in half-life during the euro period. Thus, there must be another factor working with the elimination of exchange rate shocks to offset the effect of losing the nominal exchange rate as an adjustment mechanism.

To this end, we next consider the role of changes in price dynamics, specifically the response of prices to PPP deviations. Observe that the parameter ρ_{20} in the VECM system (4) measures the equilibrium response of Δp_t to PPP deviations, i.e., the speed of mean-reversion of *p*. Analogously, the parameter γ_0 in equation (5) measures the speed of mean reversion of relative prices estimated during the euro period, with a higher absolute value of γ_0 indicating faster mean-reversion of prices and hence of the real exchange rate during the euro period. Inspection of Tables 2 and 3 indicates that the estimated value of ρ_{20} during the pre-euro period (-0.01) is smaller in absolute value than that of γ_0 in equation (5) during the euro period (-0.040). This indicates that price adjustment became faster after the introduction of the euro. This is consistent with claims that the introduction of a common currency promotes price transparency and arbitrage.[19] It also suggests a reason why the half-life of *q* fell in response to *p* shocks during the euro period.

To assess the quantitative impact of increasing the long-run dynamic response of *p* to PPP deviations, in simulation 5 we run an experiment that increases the absolute value of ρ_{20} from ρ_{20} (= -0.01) to $\hat{\gamma}_0$ (= -0.040), and find that the half-life conditional on both exchange rate

[19] This hypothesis is also tested in Huang and Yang (2014). We find that the elimination of exchange rate shocks is just as important.

and price shocks falls by 41% (from 2.00 in the benchmark to 1.19). This is about the same amount by which the half-life of *q* fell in simulation 1 when eliminating exchange rate shocks and only allowing price shocks.

We conclude that in isolation each of these changes – the elimination of *e* as a source of shocks (simulation 1) and the strengthening of long-run price dynamics (simulation 5) -- contribute to the decline in real exchange rate persistence. Hence both factors work to offset the tendency for the half-life to rise in response to the loss of the nominal exchange rate as an adjustment mechanism (simulation 4). In simulation 6 in Table 4 we combine these three experiments together by simultaneously shutting down the long-run equilibrium adjustment of relative price changes to exchange rate changes, strengthening the long-run response equilibrium adjustment of relative prices to relative price changes, and conditioning on price shocks. Observe that in the absence of exchange rate shocks, the half-life falls to 1.15, even below that estimated for the euro period from the autoregression in Table 2 (1.61). Thus, even though losing the exchange rate as an adjustment mechanism can dramatically amplify the half-life (10.65 in simulation 4 of Table 4), this is more than offset by the faster adjustment created by the combination of eliminating the exchange rate as a source of shocks along with a greater long-run dynamic price response.

We also ran a number of other experiments changing the remaining parameters in various combinations, and did not find any cases with a large effect on the half-life. We conclude that the three effects identified above are the key drivers of the decline in real exchange rate half-life after the introduction of the euro. The loss of the exchange rate as an adjustment mechanism was more than compensated by the elimination of the exchange rate as a source of shocks, in combination with faster price level adjustment.

5. Conclusions

While economic theory has highlighted the usefulness of flexible exchange rates in promoting adjustment of international relative prices, flexible exchange rates also can be a source of destabilizing shocks leading to large and persistent relative price deviations. Our study is motivated by the finding that when countries joining the euro currency union abandoned their national exchange rates, the speed of equilibrium real exchange rate adjustment increased, implying deviations from purchasing power parity (PPP) were eliminated more quickly. This finding lends support to recent claims that flexible nominal exchange rates are not essential to the promotion of international relative price adjustment.

To disentangle the possible causes for this finding we develop a novel methodology for conducting counterfactual simulations of an estimated VECM that distinguishes between the roles of the nominal exchange rate as an adjustment mechanism and as a source of shocks. We find evidence that prior to adoption of the euro these countries relied upon nominal currency adjustment as a mechanism to correct for PPP deviations arising from divergent domestic inflation rates. However, the loss of the exchange rate as an adjustment mechanism was more than compensated by the elimination of the exchange rate as a source of shocks, in combination with faster price level adjustment after the introduction of the euro.

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Table 1. Nested Autoregression Estimates and Half-life of Real Exchange Rate

	Pre-euro period		Euro period		$\eta_{1,m}-\eta_{2,m}$	
$q_{j,t-1}$	1.153**	(1.111, 1.196)	0.919**	(0.854, 0.985)	0.235**	(0.153, 0.311)
$q_{j,t-2}$	-0.235**	(-0.299, -0.172)	-0.011	(-0.100, 0.076)	-0.225**	(-0.333, -0.118)
$q_{j,t-3}$	0.138**	(0.072, 0.201)	-0.013	(-0.091, 0.066)	0.150**	(0.051, 0.249)
$q_{j,t-4}$	-0.202**	(-0.267, -0.136)	0.100**	(0.027, 0.170)	-0.302**	(-0.396, -0.203)
$q_{j,t-5}$	0.118**	(0.056, 0.185)	-0.145**	(-0.209, -0.079)	0.263**	(0.172, 0.352)
$q_{j,t-6}$	0.121**	(0.053, 0.184)	0.414**	(0.347, 0.476)	-0.294**	(-0.389, -0.200)
$q_{j,t-7}$	-0.188**	(-0.252, -0.123)	-0.391**	(-0.458, -0.321)	0.203**	(0.110, 0.297)
$q_{j,t-8}$	0.046	(-0.018, 0.108)	0.156**	(0.088, 0.226)	-0.110**	(-0.205, -0.020)
$q_{j,t-9}$	0.129**	(0.062, 0.193)	-0.143**	(-0.209, -0.077)	0.272**	(0.178, 0.365)
$q_{j,t-10}$	-0.084**	(-0.146, -0.020)	0.066**	(0.019, 0.114)	-0.150**	(-0.228, -0.070)
$q_{j,t-11}$	-0.023	(-0.066, 0.017)	----		----	
					dHL=*HL*1-*HL*2	
Half-life	2.393	(1.808, 3.689)	1.348	(1.028, 2.013)	1.045*	(0.125, 2.296)

Note: Table reports estimates for the equation

$$q_{j,t} = d_{pre-euro,t}(a_{1j} + b_{1j}t + \sum_{m=1}^{M_1}\eta_{1m}q_{j,t-m}) + (d_{euro,t})(a_{2j} + b_{2j}t + \sum_{m=1}^{M_2}\eta_{2m}q_{j,t-m}) + \varepsilon_{j,t}$$

augmented with the cross-sectional means of the dependent and explanatory variables. Coefficients are estimated using the common correlated effects pooled (CCEP) methodology of Pesaran (2006) and bias adjusted using Kilian (1998) double bootstrap method with 1000 iterations. $d_{pre-euro,t}$ ($d_{euro,t}$) is a regime dummy variable that takes a value of 1 (0) for years during the pre-euro period and a value of 0 (1) for the euro period. Numbers in parentheses are 5% and 95% confidence intervals of estimates constructed from the double bootstrap method of Kilian with 2000 iterations. ** and * indicate statistical significance at the 5% and 10% level, respectively. Half-lives in years are calculated from the simulated impulse response function derived from parameter estimates. *dHL* is the difference in half-lives between the pre-euro (*HL*1) and euro periods (*HL*2).

Table 2. Non-nested Autoregression Estimates and Half-life of Real Exchange Rate

	Pre-euro period		Euro period (q=p)	
$q_{j,t-1}$	1.154**	(1.112, 1.197)	0.944**	(0.890, 0.998)
$q_{j,t-2}$	-0.237**	(-0.300, -0.173)	-0.027	(-0.104, 0.044)
$q_{j,t-3}$	0.138**	(0.072, 0.202)	-0.013	(-0.085, 0.061)
$q_{j,t-4}$	-0.201**	(-0.267, -0.135)	0.071*	(0.004, 0.142)
$q_{j,t-5}$	0.119**	(0.056, 0.185)	-0.070*	(-0.136, -0.001)
$q_{j,t-6}$	0.121**	(0.054, 0.185)	0.406**	(0.339, 0.469)
$q_{j,t-7}$	-0.189**	(-0.253, -0.124)	-0.417**	(-0.484, -0.346)
$q_{j,t-8}$	0.047	(-0.017, 0.109)	0.090**	(0.019, 0.162)
$q_{j,t-9}$	0.129**	(0.063, 0193)	-0.076*	(-0.149, -0.003)
$q_{j,t-10}$	-0.085**	(-0.146, -0.020)	0.052	(-0.002, 0.106)
$q_{j,t-11}$	-0.023	(-0.066, 0.017)	----	
Half-life	2.392	(1.807, 3.687)	1.610	(1.170, 2.671)

Note: Table reports estimates for the equations

Pre-euro: $q_{j,t} = a_{1,j} + b_{1,j}t + \sum_{m=1}^{M_1} \eta_{1m} q_{j,t-m} + \varepsilon_{j,t}, \;\; j = 1,...,N, \text{ and } t = 1,...,T_1.$

Euro: $p_{j,t} = a_{2,j} + b_{2,j}t + \sum_{m=1}^{M_2} \eta_{2m} p_{j,t-m} + \varepsilon_{j,t}, \;\; j = 1,...,N, \text{ and } t = T_1 + 1 + M_2,...,T.$

augmented with the cross-sectional means of the dependent and explanatory variables. Coefficients are estimated using common correlated effects pooled (CCEP) methodology of Pesaran (2006) and bias adjusted using Kilian (1998) double bootstrap method with 1000 iterations. Numbers in parentheses are 5% and 95% confidence intervals constructed from the double bootstrap method of Kilian with 2000 iterations. ** and * indicate statistical significance at the 5% and 10% level, respectively. Since the nominal exchange rate is fixed during the euro period, $q_{j,t} = p_{j,t}$ during this period. Half-life of real exchange rate q in years is calculated from the simulated impulse response function derived from parameter estimates.

Table 3. VECM Estimates and Half-life of Real Exchange Rate for Pre-Euro Period

	Δe_t equation		Δp_t equation	
$q_{j,t-1}$	-0.022**	(-0.033, -0.013)	-0.010**	(-0.015, -0.004)
$\Delta e_{j,t-1}$	0.259**	(0.217, 0.303)	-0.033**	(-0.057, -0.010)
$\Delta e_{j,t-2}$	-0.056**	(-0.100, -0.009)	-0.008	(-0.032, 0.016)
$\Delta e_{j,t-3}$	0.068**	(0.024, 0.111)	-0.016	(-0.040, 0.010)
$\Delta e_{j,t-4}$	-0.082**	(-0.125, -0.039)	0.029**	(0.006, 0.054)
$\Delta e_{j,t-5}$	0.061**	(0.018, 0.105)	-0.024	(-0.048, 0.001)
$\Delta e_{j,t-6}$	0.024	(-0.019, 0.067)	0.012	(-0.013, 0.037)
$\Delta e_{j,t-7}$	-0.017	(-0.062, 0.028)	-0.015	(-0.040, 0.008)
$\Delta e_{j,t-8}$	0.017	(-0.027, 0.061)	-0.008	(-0.032, 0.016)
$\Delta e_{j,t-9}$	0.080**	(0.036, 0.124)	0.005	(-0.019, 0.028)
$\Delta p_{j,t-1}$	0.087*	(0.012, 0.161)	0.039	(-0.012, 0.088)
$\Delta p_{j,t-2}$	0.098**	(0.027, 0.170)	-0.046	(-0.095, 0.005)
$\Delta p_{j,t-3}$	-0.059	(-0.134, 0.011)	0.080**	(0.033,0.130)
$\Delta p_{j,t-4}$	-0.056	(-0.132, 0.020)	-0.075**	(-0.125, -0.026)
$\Delta p_{j,t-5}$	-0.040	(-0.120, 0.033)	-0.061**	(-0.108, -0.012)
$\Delta p_{j,t-6}$	-0.099**	(-0.178, -0.026)	0.127**	(0.079, 0.175)
$\Delta p_{j,t-7}$	-0.124**	(-0.201, -0.047)	-0.039	(-0.089, 0.009)
$\Delta p_{j,t-8}$	0.096**	(0.020, 0.170)	-0.067**	(-0.113, -0.019)
$\Delta p_{j,t-9}$	0.003	(-0.073, 0.080)	0.082**	(0.034, 0.130)
e shock		Half-life of *q* = 2.332, (1.738, 3.379)		
p shock		Half-life of *q* = 1.141, (0.518, 2.168)		
e, *p* shocks together		Half-life of *q* = 2.003, (1.481, 2.874)		

Note: Table reports estimates for the system

$$\begin{bmatrix}\Delta e_{j,t}\\ \Delta p_{j,t}\end{bmatrix}=\begin{bmatrix}g_{10,j}\\ g_{20,j}\end{bmatrix}+\begin{bmatrix}g_{11,j}\\ g_{21,j}\end{bmatrix}t+\begin{bmatrix}\rho_{10}\\ \rho_{20}\end{bmatrix}q_{jt-1}+\begin{bmatrix}a_{11} & b_{11}\\ c_{11} & d_{11}\end{bmatrix}\begin{bmatrix}\Delta e_{jt-1}\\ \Delta p_{jt-1}\end{bmatrix}+....+\begin{bmatrix}a_{1M-1} & b_{1M-1}\\ c_{1M-1} & d_{1M-1}\end{bmatrix}\begin{bmatrix}\Delta e_{j,t-M+1}\\ \Delta p_{j,t-M+1}\end{bmatrix}+\begin{bmatrix}\varepsilon_{ej,t}\\ \varepsilon_{pj,t}\end{bmatrix}$$

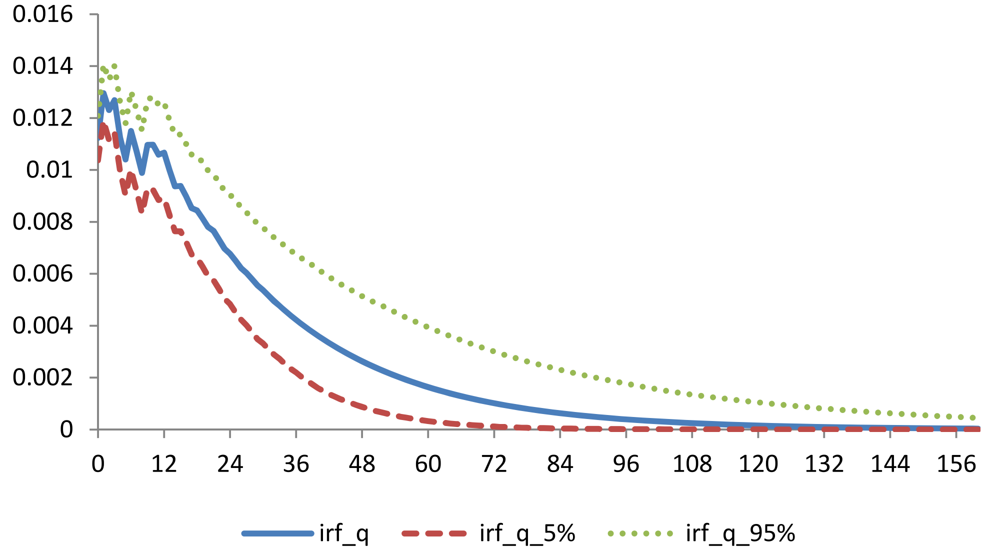
augmented with the cross-sectional means of the dependent and explanatory variables. Coefficients are estimated using common correlated effects pooled (CCEP) methodology of Pesaran (2006) and bias adjusted using Kilian (1998) double bootstrap method with 1000 iterations. Numbers in parentheses are 5% and 95% confidence intervals constructed from the double bootstrap method of Kilian with 2000 iterations. ** and * indicate statistical significance at the 5% and 10% level, respectively. Half-lives of real exchange rate conditional on shocks are reported in years and are calculated from the simulated impulse response function derived from parameter estimates.

Table 4. Counterfactual Simulations

Simulation	Half-life of Real Exchange Rate (in years)		
	e shock	*p* shock	simultaneous *e*, *p* shocks
1. Benchmark	2.332	1.141	2.003
2. Nest AR of *p* estimated for the euro period: $\rho_{10} = a_{11} = ... = a_{1k-1} = b_{11} = ... = b_{1M-1} = 0$, $c_{11} = ... = c_{1M-1} = 0$, $d_{11} = \hat{\gamma}_1$, ... $d_{1M-1} = \hat{\gamma}_{1M-1}$, $\rho_{20} = \hat{\gamma}_0$.	1.1495	1.603	1.302
3. Remove short-run response to Δe_{t-m} in Δe_t equation: $a_{11} = ... = a_{1M-1} = 0$.	1.714	1.639	1.695
4. Remove long-run response to q_{t-1} in Δe_t equation: $\rho_{10} = 0$.	10.648	5.382	9.180
5. Strengthen long-run response to q_{t-1} in Δp_t equation: $\rho_{20} = \hat{\gamma}_0$.	1.444	0.536	1.188
6. Remove long-run response to q_{t-1} in Δe_t equation and strengthen long-run response to q_{t-1} in Δp_t equation: $\rho_{20} = \hat{\gamma}_0$, $\rho_{10} = 0$.	2.652	1.151	2.222

Note: Table reports counterfactualsimulations based on VECM estimates reported for equation (4) for the pre-euro period in Table 3. Simulations 2, 5, and 6 make use of non-nested AR estimates of equation (3.2) for *p* reported in Table 2 for the euro period in order to recover estimates of the γ_i coefficients in equation (5), indicated by hats (^), as discussed in the text. Half-lives conditional on individual shocks are reported in years and are calculated from simulated impulse response function derived from restricted parameter values in each simulation.

A. *q* response during pre-euro period



B. *q* response during euro period

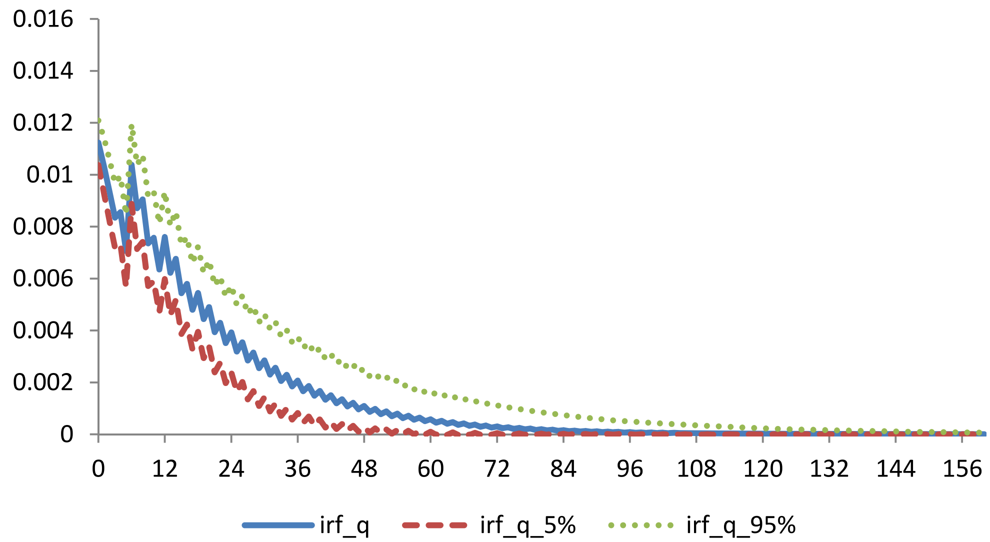
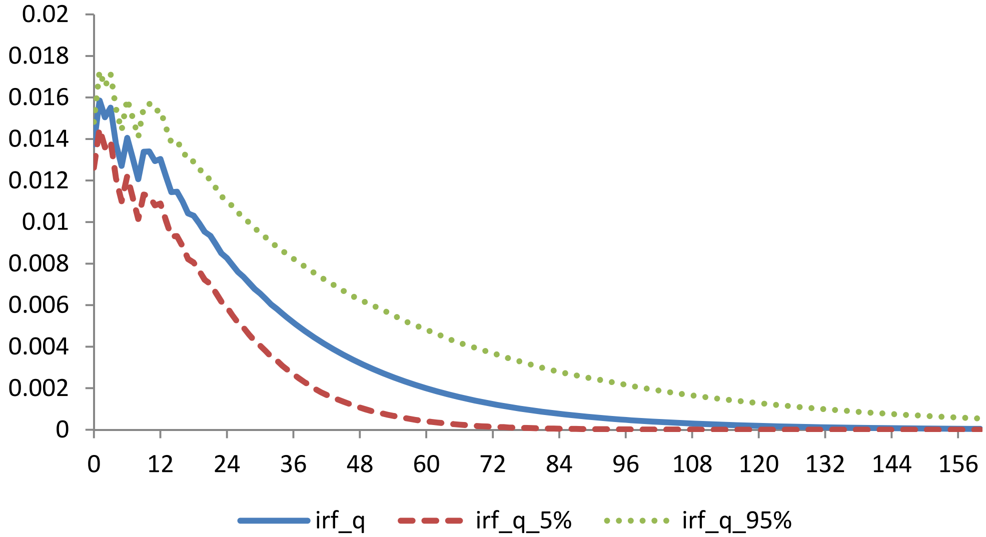


Figure 1. The impulse response function (IRF) in months of the real exchange rate to a one standard-deviation shock during the pre-euro and euro periods, respectively, based on the bias-corrected CCEP estimates of the autoregression equation (2) reported in Table 1. Dashed lines are 5% and 95% confidence intervals, constructed using the double bootstrap method of Kilian (1998) with 2000 iterations.

A. *q* response during pre-euro period



B. *p* response during euro period

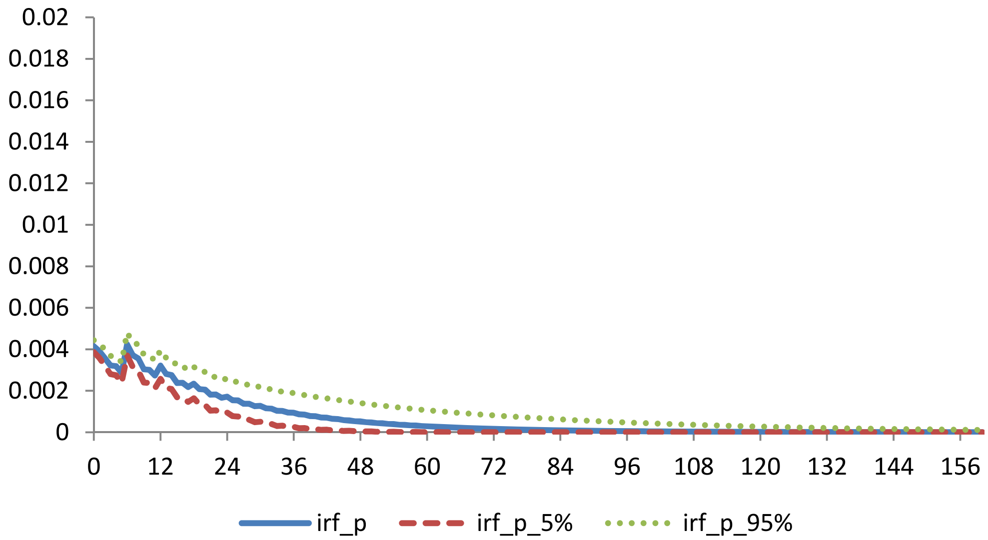
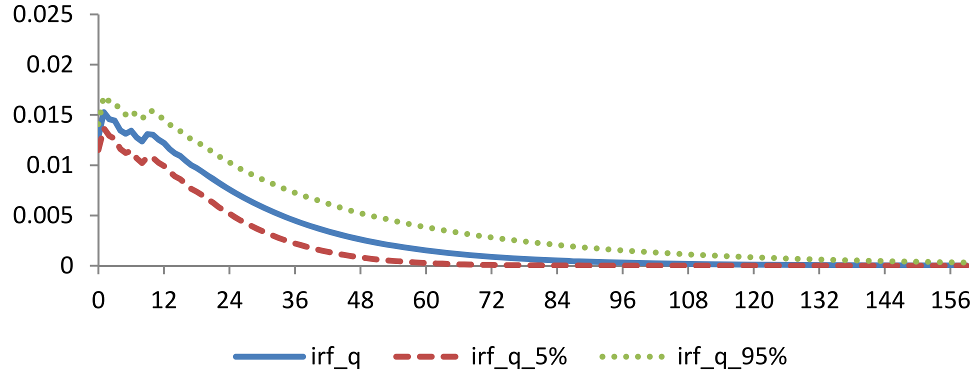
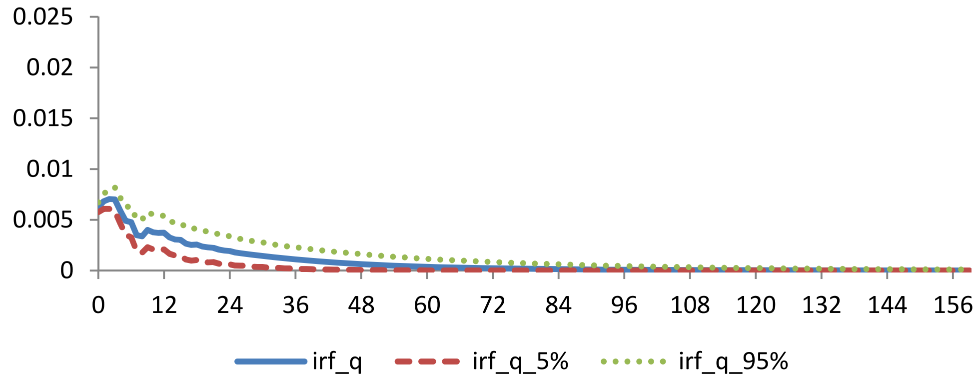


Figure 2. The impulse response function (IRF) in months of the real exchange rate *q* during the pre-euro period and of the relative price *p* during the euro period to a one standard-deviation shock, based on bias-corrected CCEP estimates of the autoregression equations (3.1) and (3.2) reported in Table 2. Dashed lines are 5% and 95% confidence intervals, constructed using the double bootstrap method of Kilian (1998) with 2000 iterations.

A. *q* response to *e* shock during pre-euro period



B. *q* response to *p* shock during pre-euro period



C. *q* response to simultaneous *e* and *p* shocks during pre-euro period

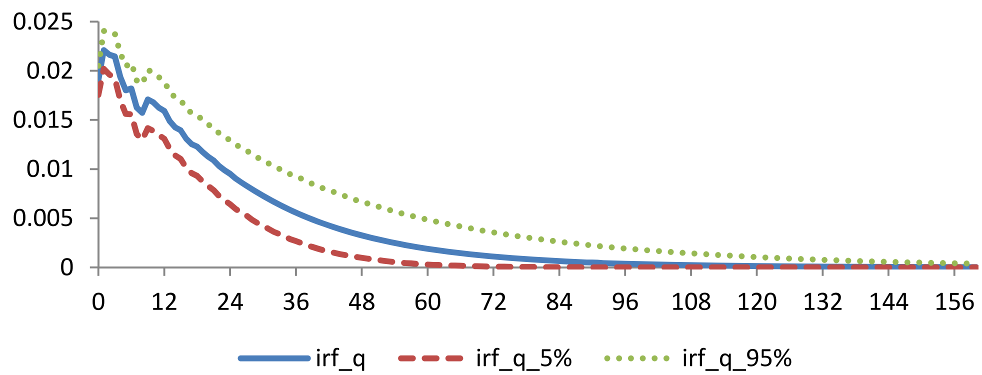
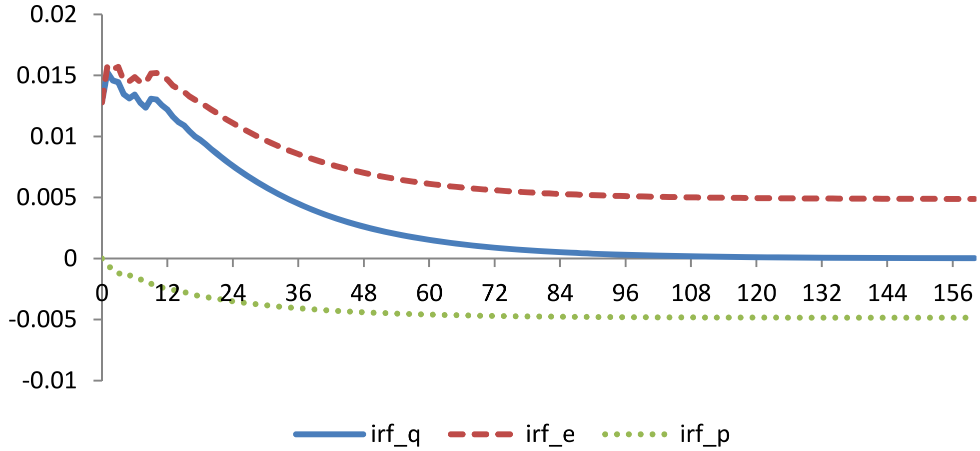


Figure 3. The impulse response function (IRF) of the real exchange rate *q* in months conditional on one standard-deviation shocks of the exchange rate, prices, and both variables simultaneously , respectively, during the pre-euro period. Dashed lines are 5% and 95% confidence intervals, constructed using the double bootstrap method of Kilian (1998) with 2000 iterations.

A. *e*, *p*, and *q* response to *e* shock during pre-euro period



B. *e*, *p*, and *q* response to *p* shock during pre-euro period

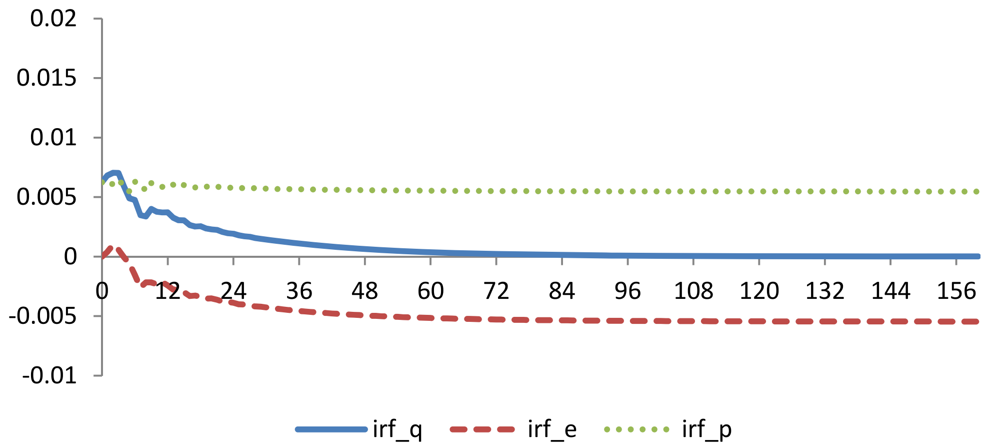


Figure 4. The impulse response function (IRF) of the nominal exchange rate, the real exchange rate, and the price level in months conditional on one standard-deviation shocks of the nominal exchange rate and prices, respectively, during the pre-euro period based on bias-corrected CCEP estimates of the VECM of equation (4) reported in Table 3.