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"Conditional PPP" and real exchange rate convergence in the euro area

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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 their long-run equilibrium surprisingly became faster. To investigate, we distinguish between differing rates of purchasing power parity (PPP) convergence conditional on alternative shocks, which we refer to as "conditional PPP." We find that 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 prices. These findings support claims that flexible exchange rates are not necessary to promote long-run international relative price adjustment.

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,

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Berka et al. (2012) argued that the real exchange rate adjustment in a currency union like the euro area might be faster than 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).[1] In addition to providing evidence regarding how the euro has affected market integration, this investigation also provides two broader lessons regarding how to understand real exchange rate dynamics. First, we distinguish between PPP convergence conditional on alternative shocks, which we refer to as "conditional PPP." Since we show that the half-life of real exchange rate convergence can differ significantly depending on what was the source of the real exchange rate disturbance, we argue it is important for international macroeconomists to make this distinction when characterizing the relevance of PPP as a theory of real exchange rate behavior. PPP may hold in the context of some shocks while not holding well for others, so whether PPP is a useful characterization of a given country or period depends on the mix of shocks prevalent in that sample. Second, we distinguish between the roles of the nominal exchange rate as source of shocks and as a mechanism of adjustment to shocks.

We develop a stochastic simulation-based methodology to examine these two key distinctions in characterizing real exchange rate dynamics. 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. Our exchange rate shocks could be viewed as shocks to asset preferences in an interest rate parity or portfolio preference relation, as in the framework of Flood and Rose (1999). We next estimate the half-life of the real exchange rate adjustment conditional on specific shocks, which is where "conditional PPP" comes to the fore. 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.

Our estimations allow for a linear trend in real exchange rates, following the practice of Taylor (2002) and Papell and Prodan (2006). This has been motivated theoretically in terms of Balassa-Samuelson effects of productivity differentials between traded and nontraded sectors. Given the wide belief that some European countries have experienced productivity catchup and corresponding Balassa-Samuelson effects as a result of greater integration (see Canzoneri et al., 2002; and Berka et al., 2014), allowing for a deterministic trend seems especially appropriate for this dataset.

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 relied 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 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-life 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-life 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 et al. (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 of monetary regime shifts, such as the adoption of the euro, on real exchange rate persistence. Parsley and Popper (2001) find faster real exchange rate convergence under currency pegs, but they do not study the case of the euro common currency. While they study in detail the nominal exchange rate as an adjustment mechanism, they do not examine the competing role of the nominal exchange rate as a source of shocks.

Several papers have investigated PPP adjustment during the euro period. Koedijk et al. (2004), Lopez and Papell (2007), and Zhou et al. (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 (2015) 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.

[1] There is a large and long-standing literature estimating rates of convergence to PPP. See Imbs et al. (2005) for a prominent example and discussion of this literature.

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.

Earlier work in Bergin et al. (2014) shares elements of our VECM methodology when studying the change in real exchange rate volatility and persistence during the transition from Bretton Woods to post-Bretton Woods periods, focusing on the implications for the appropriateness of sticky price models. However, that paper does not study the competing roles of the nominal exchange rate as a source of shocks as well as an adjustment mechanism to shocks, which is the primary question addressed in the present paper. Using the transition to the euro as a natural and clean experiment, the present paper finds that the distinction between shocks is essential to explaining the change in half-life of the real exchange rate between periods. Thus, the present paper is important as a platform from which we can introduce "conditional PPP" as an important new concept in international macroeconomics.

Our work is complementary to, but distinct from, the literature studying price dispersion in the euro area using micro level data on individual goods prices. Recent work in this area tends to find that introduction of the euro reduced the degree of price dispersion, suggesting increased integration of national goods markets within the euro area (see Glushenkova and Zachariadis, 2016). This conclusion is consistent with our finding of faster real exchange rate convergence after euro adoption. However, study of cross-section dispersion in micro level data on individual goods prices is distinct from our study of the dynamic adjustment of price aggregates, both in methodology and purpose. For example, a prominent conclusion of the micro level literature on European price dispersion is that for any given pair of countries there are roughly as many overpriced as underpriced goods (e.g., Crucini et al., 2005), suggesting that exchange rate movements cannot reduce price dispersion for all goods at the micro level at the same time. So the micro literature tends not to study the nominal exchange rate as an adjustment mechanism.

Eichenbaum et al. (2016) also present evidence that real exchange rate adjustment occurs primarily via nominal exchange rate changes under a flexible exchange rate regime, but via price changes under a fixed rate regime. This evidence supports our conclusions, but differs in that it is generated from univariate regressions of nominal exchange rate and relative price change on real exchange rates over various horizons, rather than, as we do, from a VECM, which jointly estimates exchange rate and price changes, controls for short run dynamics, and enables estimation of the overall half-life of the real exchange rate.

The paper is organized as follows. The data are presented in Section 2. The main empirical results are presented in the following two sections, with Section 3 estimating the half-life 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 discusses 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*.[2] The sample is monthly in frequency and covers the period April 1973–February 2016. 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.[3] 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) estimating the panel regression:

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

where $\bar{q}_t = \sum_{j=1}^{N} q_{j,t}$ is the cross-section mean of $q_{j,t}$ across the N country exchange rates, $\Delta\bar{q}_t = \bar{q}_t - \bar{q}_{t-1}$, and the purpose of augmenting the specification with cross-section means is to control for contemporaneous correlation among $\varepsilon_{j,t}$.

All of our estimation equations detrend the real exchange rate by including a regime-specific constant and time trend. This follows the practice of Taylor (2002) and Papell and Prodan (2006), who motivate this in terms of Balassa-Samuelson effects and argue that productivity differentials in traded goods between countries determine the domestic relative price of nontradables, which in turn lead to trend deviations from PPP. Obstfeld (1993) utilized this idea to explain why real

[2] 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 because its currency was pegged 1:1 to the Belgium franc before joining the euro area. The price data are not seasonally adjusted.

[3] This specification of the log difference in domestic price indices 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).

exchange rates should contain a deterministic trend. Papell and Prodan refer to this modified version of PPP as "Trend PPP". Given the wide belief that some European countries have experienced productivity catchup and corresponding Balassa-Samuelson effects as a result of greater integration (see Canzoneri et al., 2002; and Berka et al., 2014), allowing for a deterministic trend seems especially appropriate for this dataset.[4]

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:

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

where $t\text{-}stat_j(N,T)$ is the t statistic of $\omega_{0,j}$ from the estimation of Eq. (1). The AIC criterion is applied to select the appropriate lag order in Eq. (1). Setting the maximum lag length of M to 12, the selected optimal lag length based on median AICs is 10 ($M = 11$) for the whole period.[5] Based on the selected lagged order, the unit-root hypothesis for q is rejected at the 5% level ($t\text{-}stat = -3.14$).

As an additional diagnostic, we also test for evidence of nonlinearity of the smooth transition autoregressive (STAR) type (see Michael et al., 1997, and Taylor et al., 2001 for example), as nonlinearity could affect our estimated half-lives. We follow Granger and Terasvirta (1993) and Terasvirta (1998) by estimating the following auxiliary regression for each country j during the pre-euro and euro periods:

$$q_{j,t} = \pi_{j0} + \sum_{m=1}^{M_j}\left(\pi^1_{jm}q_{j,t-m} + \pi^2_{jm}q_{j,t-m}q_{j,t-d_j} + \pi^3_{jm}q_{j,t-m}q^2_{j,t-d_j} + \pi^4_{jm}q_{j,t-m}q^3_{j,t-d_j}\right) + \varepsilon_{j,t}, \quad j = 1,\ldots,N$$

where $q_{j,t-d_j}$ is the transition variable, and the optimal lag order M_j and d_j are determined from the data. The null hypothesis of linearity $H_0 : \pi^2_{jm} = \pi^3_{jm} = \pi^4_{jm} = 0 \;\; (m = 1,\cdots,M_j)$ is used to test against STAR nonlinearity with the F statistic. The results indicate that the hypothesis of linearity is rejected for only 1 out of our 9 countries (France), at the 5% level, in the pre-euro period, and it is not rejected for any country in the euro period. We conclude that there is no evidence for the importance of STAR nonlinearities for real exchange rates in our particular sample. More details are provided in the appendix in Supplementary Material.

3. Estimating rates of convergence

We begin by documenting the change in half-life of real exchange rate convergence for the euro area, estimating an autoregressive panel model.

3.1. Nested estimation

To permit significance tests for a change in half-life, we begin with a model that nests together the data for the pre-euro and euro periods:

$$q_{j,t} = d_{pre\text{-}euro,t}\left(\sum_{m=1}^{M_1}\eta_{1m}q_{j,t-m}\right) + d_{euro,t}\left(\sum_{m=1}^{M_2}\eta_{2m}q_{j,t-m}\right) + \varepsilon_{j,t}. \tag{2}$$

where the indicator regime variable $d_{pre\text{-}euro,t}$ takes a value of 1 during the pre-euro period and a value of 0 otherwise, i.e., $d_{pre\text{-}euro,t} = 1$ for $t = 1,\ldots,T_1$, the end date of the pre-euro period, and correspondingly $d_{euro,t} = 1$ for $t = T_1 + M_2 + 1,\ldots,T$ and 0 otherwise.[6,7] To control for contemporaneous correlation of residuals, the common correlated effects pooled (CCEP) regressor of Pesaran (2006) is used, involving augmentation of Eq. (2) with the cross-sectional means of dependent and explanatory variables during the two regimes.[8] 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

[4] Estimation results for the model with a regime-specific constant but without a time trend are available upon request. The optimal lag order of this model based on the median AICs is 10 for the pre-euro period and 11 for the euro period. The 5–95% confidence band is (2.38, 6.78) for the pre-euro period and (2.22, ∞) for the euro period, while the 5–95% band for the half-life differential between regimes is (-2.36, ∞). The very wide confidence band in this case precludes any conclusions.

[5] We apply 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 11,7, 8, 12, 11, 12, 10, 12, 11 months.

[6] 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 $t = T_1 + 1 + M_2$.

[7] This estimation effectively detrends the real exchange rate by including a regime-specific constant and time trend in the equation. This allows the constant and trend to differ between regimes, which follows the specifications used in Taylor (2002), a fact which we confirmed with the author.

[8] 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.

estimates for each sub-period.[9,10] The optimal lag length in Eq. (2) determined based on median AICs is $M_1 = 11$ for the pre-euro period and $M_2 = 10$ for the euro period.[11]

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 1.81–3.68 years); that for the euro period is 1.50 years (with a band of 1.05–2.04 years). This represents a 37% drop in persistence in the euro period.[13] As a test of significance, our stochastic simulations also compute the difference in half-life (0.89 years) between the two regimes; these results are reported in the last column. The 5–95% confidence band for this difference of 0.10–2.25 years excludes zero. These estimates support the conclusion that the half-life of the real exchange rate is lower in the euro period.

Fig. 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.

3.2. Non-nested estimation

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:

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

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

where $M_1 = 11$ for the pre-euro period, and $M_2 = 10$ for the euro period.[15] 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 period. We estimate the above equations along with cross-sectional means of the left-hand and right-hand variables $(\bar{q}_t, \bar{q}_{t-1}, \ldots, \bar{q}_{t-M})$ for (3.1) and $(\bar{p}_t, \bar{p}_{t-1}, \ldots, \bar{p}_{t-M})$ for (3.2). Compared to the specification in Eq. (2), Eqs. (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 Eq. (3.2), are useful in simulation exercises described in Section 4.[16]

Table 2 reports parameters of the AR(11) for the pre-euro period in the first column. The estimated half-life of the real exchange rate is 2.390 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 1.689 years, which is somewhat higher than that estimated from the AR nesting both periods together. Fig. 2 plots

[9] See the appendix of Bergin et al. (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.

[10] In other words, we make use of the mean-unbiased estimator of Kilian (1998) to estimate the parameters in (2). As discussed in Murray and Papell (2002), this mean-unbiased estimator yields results comparable to those using median unbiased methods, as both appear to be effective at reducing the bias in impulse response estimates.

[11] The AIC lag lengths for Austria, Belgium, Finland, France, Ireland, Italy, Netherlands, Portugal, and Spain for the pre-euro period are 12, 7, 2, 11, 10, 7, 11, 11, 11 months, respectively, with a median of 11 for the pre-euro period and are 11, 12, 8, 7, 9, 12, 12, 7, 10 months, respectively, with a median of 10, for the euro period.

[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.

[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.

[15] As above, all our estimations detrend the real exchange rate by including a regime-specific constant and time trend in the equation.

[16] We also estimated versions of (3.1) and (3.2) that allowed for a constant without a deterministic trend. With an optimal lag structure of 10 for the pre-euro period and 11 for the euro period, the 5–95% band for the half-life was (2.39, 6.77) for the pre-euro period, and (2.72, ∞) for the euro period. Again, the result of a very wide confidence band precludes any conclusions.

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, 1.195)	0.913**	(0.857, 0.968)	0.241**	(0.171, 0.312)
$q_{j,t-2}$	−0.235**	(−0.299, −0.171)	0.005	(−0.070, 0.080)	−0.240**	(−0.336, −0.141)
$q_{j,t-3}$	0.137**	(0.072, 0.201)	−0.024	(−0.095, 0.049)	0.161**	(0.067, 0.258)
$q_{j,t-4}$	−0.202**	(−0.266, −0.136)	0.073*	(0.005, 0.139)	−0.275**	(−0.371, −0.181)
$q_{j,t-5}$	0.118**	(0.056, 0.185)	−0.105**	(−0.166, −0.043)	0.223**	(0.135, 0.319)
$q_{j,t-6}$	0.120**	(0.053, 0.184)	0.433**	(0.371, 0.489)	−0.312**	(−0.403, −0.226)
$q_{j,t-7}$	−0.187**	(−0.251, −0.122)	−0.410**	(−0.470, −0.347)	0.223**	(0.134, 0.314)
$q_{j,t-8}$	0.046	(−0.018, 0.107)	0.105**	(0.043, 0.168)	−0.060	(−0.149, 0.029)
$q_{j,t-9}$	0.129**	(0.062, 0.193)	−0.088**	(−0.151, −0.028)	0.217**	(0.128, 0.307)
$q_{j,t-10}$	−0.085**	(−0.146, −0.020)	0.054*	(0.007, 0.099)	−0.138**	(−0.215, −0.057)
$q_{j,t-11}$	−0.023	(−0.066, 0.017)	–	–	–	–
					dHL = HL1 – HL2	
Half-life	2.392	(1.809, 3.682)	1.502	(1.053, 2.040)	0.890*	(0.101, 2.245)

Note: Table reports estimates for the equation $q_{j,t} = d_{pre\text{-}euro,t}\sum_{m=1}^{M_1}\eta_{1m}q_{j,t-m} + d_{euro,t}\sum_{m=1}^{M_2}\eta_{2m}q_{j,t-m} + \varepsilon_{j,t}$ augmented with a regime-specific constant and time trend as well as 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. $d_{pre\text{-}euro,t}$ ($d_{euro,t}$) 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).

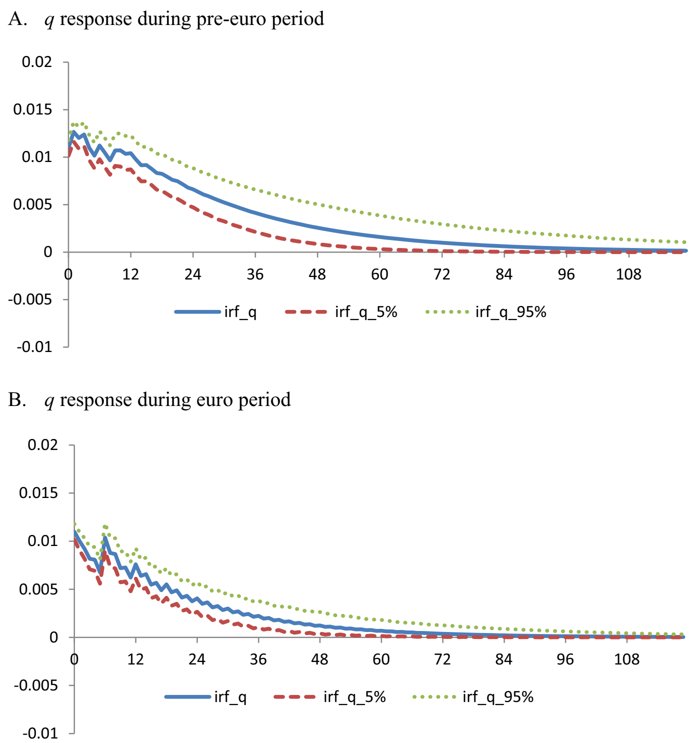

Fig. 1. The impulse response function (IRF) in months of the real exchange rate to a one standard-deviation shock during the pre-euro and euro periods, respectively, based on the bias-corrected CCEP estimates of the autoregression equation (2) reported in Table 1. Dashed lines are 5% and 95% confidence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

the estimated IRF of the real exchange rate during the pre- and euro periods, respectively; the dynamics generally appear very similar to those in Fig. 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 Fig. 1 implicitly imposes the same standard deviation for q shocks (0.011), during the pre-euro and euro periods. In contrast, the non-nested case considered in Fig. 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 0.014).

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, 1.196)	0.930**	(0.881, 0.980)
$q_{j,t-2}$	−0.237**	(−0.300, −0.173)	−0.007	(−0.075, 0.056)
$q_{j,t-3}$	0.137**	(0.071, 0.201)	−0.027	(−0.093, 0.039)
$q_{j,t-4}$	−0.201**	(−0.267, −0.135)	0.058	(−0.006, 0.124)
$q_{j,t-5}$	0.119**	(0.056, 0.185)	−0.047	(−0.107, 0.013)
$q_{j,t-6}$	0.121**	(0.054, 0.185)	0.418**	(0.357, 0.475)
$q_{j,t-7}$	−0.189**	(−0.252, −0.124)	−0.418**	(−0.482, −0.352)
$q_{j,t-8}$	0.046	(−0.017, 0.108)	0.045	(−0.021, 0.113)
$q_{j,t-9}$	0.130**	(0.063, 0.193)	−0.034	(−0.104, 0.038)
$q_{j,t-10}$	−0.085**	(−0.146, −0.020)	0.044	(−0.007, 0.095)
$q_{j,t-11}$	−0.023	(−0.066, 0.017)	–	
Half-life	2.390	(1.807, 3.688)	1.689	(1.189, 2.844)

Note: Table reports estimates for the equations $\begin{matrix}\text{Pre-euro}: q_{j,t} = \sum_{m=1}^{M_1}\eta_{1m}q_{j,t-m} + \varepsilon_{j,t}, j = 1,\ldots,N, \text{ and } t = 1,\ldots,T_1. \\ \text{Euro}: p_{j,t} = \sum_{m=1}^{M_2}\eta_{2m}p_{j,t-m} + \varepsilon_{j,t}, j = 1,\ldots,N, \text{ and } t = T_1 + 1 + M_2,\ldots,T.\end{matrix}$ augmented with a regime-specific constant and time trend as well as 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.

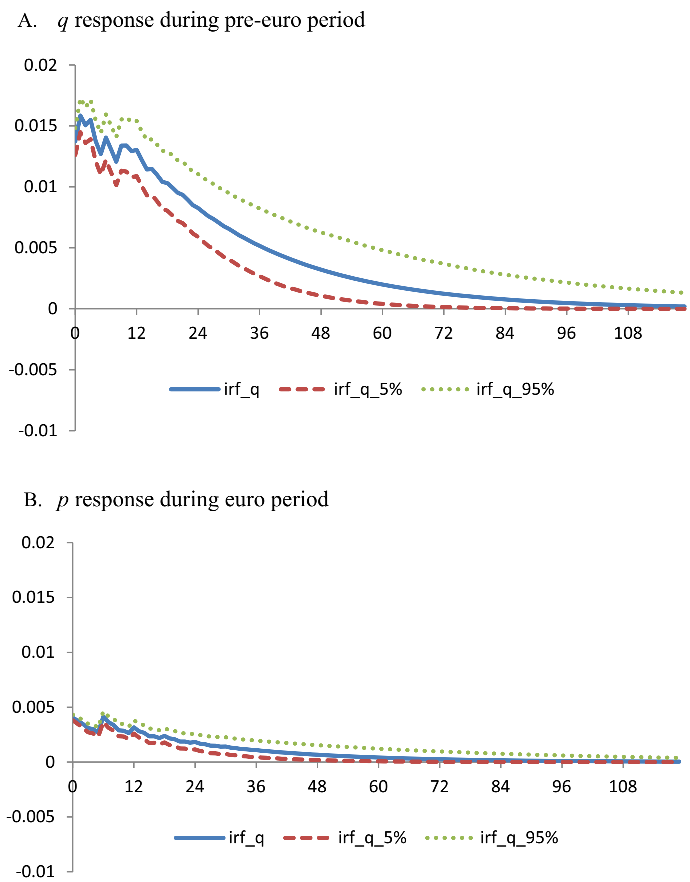

Fig. 2. The impulse response function (IRF) in months of the real exchange rate q during the pre-euro period and of the relative price p during the euro period to a one standard-deviation shock, based on bias-corrected CCEP estimates of the autoregression equations (3.1) and (3.2) reported in Table 2.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

3.3. Estimation for a comparison group

Although we associate the fall in half-life for euro area countries after 1999 with the introduction of the euro, it is possible that some other change was responsible that happened to coincide with the euro introduction. As a check, we also apply our analysis to a set of European comparison countries that did not participate in the eurozone currency arrangement. We consider a set of five countries: Denmark, Norway, Sweden, Switzerland, and the United Kingdom. If adoption of the common currency is directly associated with the fall in half-life, we would expect no significant change in half-lives between the pre-euro and euro periods for this group of countries.

As a preliminary, we confirm that we can reject a unit root for this sample. The CIPS test is −2.912 for the full period, implying that the I(1) hypothesis is rejected at the 10% level. The results from estimation of the nested model, Eq. (2), are reported in panel A of Table 3. The selected lag orders are 3 and 1 for the pre-euro and euro periods, respectively, and the estimated coefficients are all significant at the 5% level. The estimated half-life is 1.79 for the pre-euro period and 2.74 for the euro period, which clearly contrasts with the finding reported in Table 2 that the half-life for Eurozone countries fell in the later period. The half-life difference between two sub-periods is insignificant at the 10% level for these comparison countries.

We also consider a comparison group excluding Denmark. Even though Denmark did not formally join the Eurozone, it has pursued a policy of effectively pegging its currency to the euro since 1999. Thus to support our claim that exchange rate shocks are an important source of real exchange rate persistence, it seems sensible to limit our comparison group to those countries that had flexible exchange rates relative to the euro during the euro period.[17] As reported in panel B of Table 3, the estimated half-life is 1.88 years for the pre-euro period and 2.19 years for the euro period, and the half-life difference between two sub-periods is insignificant at the 10% level. We take these results as supporting our claim that the fall in half-live for euro area countries after 1999 is attributable to the introduction of the euro.

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.[18]

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:

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

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.[19] 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 $\Delta p_{j,t}$. The bias-adjusted CCEP estimates are then constructed based on Kilian (1998). An optimal lag length of $M = 10$ is determined from the median of AICs of vector autoregressions of $e_{j,t}$ and $p_{j,t}$ for individual countries over the pre-euro period.[20]

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:

$$\Delta p_{j,t} = \gamma_0 p_{j,t-1} + \gamma_1 \Delta p_{j,t-1} + \gamma_2 \Delta p_{j,t-2} + \cdots + \gamma_{M_2-1} \Delta p_{j,t-M_2+1} + \varepsilon_{pj,t}. \tag{5}$$

[17] This is the same comparison considered by Huang and Yang (2015).

[18] We employ this methodology in Bergin et al. (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.

[19] As above, all our estimations detrend the real exchange rate by including a regime-specific constant term and time trend.

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

Table 3
Nested autoregression estimates and half-life of real exchange rate for non-eurozone countries.

	Pre-euro period		Euro period		$\eta_{1,m} - \eta_{2,m}$	
A. 5 non-eurozone countries: Denmark, Norway, Sweden, Switzerland, United Kingdom						
$q_{j,t-1}$	1.278**	(1.228, 1.332)	0.979**	(0.959, 0.995)	0.299**	(0.246, 0.357)
$q_{j,t-2}$	−0.400**	(−0.484, −0.322)	–	–	–	–
$q_{j,t-3}$	0.088**	(0.038, 0.142)	–	–	–	–
					$dHL = HL1 - HL2$	
Half-life	1.793	(1.253, 2.882)	2.738	(1.368, 10.138)	0.945	(−9.515, 0.778)
B. 4 non-eurozone countries: Norway, Sweden, Switzerland, United Kingdom						
$q_{j,t-1}$	1.292**	(1.236, 1.349)	1.088**	(1.016, 1.158)	0.204**	(0.115, 0.299)
$q_{j,t-2}$	−0.423**	(−0.520, −0.335)	−0.116**	(−0.183, −0.046)	−0.307**	(−0.427, −0.194)
$q_{j,t-3}$	0.098**	(0.042, 0.159)	–	–	–	–
					$dHL = HL1 - HL2$	
Half-life	1.875	(1.249, 3.244)	2.194	(1.198, 6.009)	−0.318	(−4.208, 1.316)

Note: Table reports estimates for the following equation augmented with a regime-specific constant and time trend as well as the cross-sectional means of the dependent and explanatory variables: $q_{j,t} = d_{pre\text{-}euro,t}\sum_{m=1}^{M_1}\eta_{1m}q_{j,t-m} + d_{euro,t}\sum_{m=1}^{M_2}\eta_{2m}q_{j,t-m} + \varepsilon_{j,t}$. 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. $d_{pre\text{-}euro,t}$ ($d_{euro,t}$) 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).

The coefficients in (5) can be obtained by a simple transformation of the coefficients (η_{2m}) in Eq. (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 $m = 1,2,\ldots,M_2 - 1$. 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, $\hat{\gamma}_0 = -0.038$.

Under appropriate parameter restrictions, the VECM can be seen to nest Eq. (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 4 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 ($\hat{\rho}_{10} = -0.022$, $\hat{\rho}_{20} = -0.010$) are both 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.

Our "exchange rate" shocks may be interpreted in the context of the framework of Flood and Rose (1999), where exchange rate fluctuations are driven by asset market shocks, modeled as shocks to an interest rate parity equation and/or asset preferences. Flood and Rose assume these shocks are much more volatile than fundamental macroeconomic shocks, and regard them as the dominant factor driving volatile exchange rate fluctuations when exchange rates are flexible, but which disappear under a fixed exchange rate regime. They motivate this theoretical approach with the empirical finding that exchange rates, but not macro fundamentals, are much more volatile under a floating exchange rate regime. In turn, we associate our slower moving price innovations with the more standard macro-fundamentals shocks in their framework. It is well known in the literature that there never will be an exact correspondence between the shocks estimated from VAR identification via Cholesky ordering and the shocks in standard theoretical models (see Canova and Pina, 1999). But in choosing our Cholelsky ordering with the exchange rate ordered before prices, we follow the logic of Flood and Rose, that the volatile fluctuations in exchange rates are driven by volatile exogenous shocks to the foreign exchange market, and that these shocks disappear with a fixed exchange rate regime.

Table 4 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.

Table 4
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, 0.017)
$\Delta e_{j,t-3}$	0.068**	(0.024, 0.111)	−0.016	(−0.040, 0.010)
$\Delta e_{j,t-4}$	−0.082**	(−0.125, −0.038)	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, 0.009)
$\Delta e_{j,t-8}$	0.017	(−0.027, 0.061)	−0.009	(−0.032, 0.016)
$\Delta e_{j,t-9}$	0.080**	(0.036, 0.125)	0.005	(−0.019, 0.028)
$\Delta p_{j,t-1}$	0.086*	(0.011, 0.160)	0.039	(−0.012, 0.089)
$\Delta p_{j,t-2}$	0.098**	(0.027, 0.169)	−0.046	(−0.095, 0.005)
$\Delta p_{j,t-3}$	−0.059	(−0.134, 0.011)	0.079**	(0.033, 0.129)
$\Delta p_{j,t-4}$	−0.056	(−0.132, 0.020)	−0.075**	(−0.125, −0.026)
$\Delta p_{j,t-5}$	−0.041	(−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**	(−0.200, −0.047)	−0.039	(−0.089, 0.009)
$\Delta p_{j,t-8}$	0.096**	(0.021, 0.170)	−0.068**	(−0.113, −0.019)
$\Delta p_{j,t-9}$	0.003	(−0.073, 0.080)	0.083**	(0.034, 0.130)
e shock		Half-life of *q* = 2.331, (1.739, 3.378)		
p shock		Half-life of *q* = 1.139, (0.518, 2.165)		
e, *p* shocks together		Half-life of *q* = 2.002, (1.477, 2.874)		

Note: Table reports estimates for the system $\begin{bmatrix}\Delta e_{j,t}\\ \Delta p_{j,t}\end{bmatrix} = \begin{bmatrix}\rho_{10}\\ \rho_{20}\end{bmatrix} q_{jt-1} + \begin{bmatrix}a_{11} b_{11}\\ c_{11} d_{11}\end{bmatrix}\begin{bmatrix}\Delta e_{jt-1}\\ \Delta p_{jt-1}\end{bmatrix} + \ldots + \begin{bmatrix}a_{1M-1} b_{1M-1}\\ c_{1M-1} d_{1M-1}\end{bmatrix}\begin{bmatrix}\Delta e_{j,t-M+1}\\ \Delta p_{j,t-M+1}\end{bmatrix} + \begin{bmatrix}\varepsilon_{ej,t}\\ \varepsilon_{pj,t}\end{bmatrix}$ augmented with a regime-specific constant and time trend as well as 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.

Fig. 3 plots impulse responses (IRFs) of the real exchange rate to nominal 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, Fig. 3 supports the characterization of the results reported in Table 4. For later reference, Fig. 4 plots the impulse responses for the separate components of the real exchange rate, that is, the nominal exchange rate and the price index ratio.

To help clarify where our results come from, we construct measures of exchange rate and price volatility for individual countries over time. Specifically, we compute the exchange rate and price shocks for each country implied by the estimated VECM specification in Eq. (4), and calculate standard deviations for each year, using the twelve monthly observations for the year.[21] Fig. 5 plots the resulting time series of exchange rate and price standard deviations for each country. Observe that high volatility of exchange rate shocks is not special to any one country or any one period, but is a widespread phenomenon in our sample. We infer that our results are not driven by rare or extreme cases in the data. Observe as well that exchange rate shocks were typically more volatile than price shocks for almost all periods and countries, with the exceptions of Austria and the Netherlands, who maintained exchange rate policies tied closely to Germany throughout the sample, beginning very shortly after the end of Bretton Woods.

The evident variations in exchange rate shock variability lend themselves to easy interpretation. The 1970s was a period of higher than usual exchange rate volatility for several countries (Italy, Spain, Portugal, France) due to periods of relatively high domestic inflation and less strict monetary policy than in Germany. Volatility declined for most countries during the 1980s, as they successfully coordinated their exchange rate policies as part of the European Monetary System.[22] Volatility for several countries rose again in the early 1990s (notably Italy, Ireland, Finland, Portugal, and Spain) in the aftermath of the EMS crisis of 1992.

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

[21] We compute pooled residuals from (4) using the CCEP-estimated common coefficients, with country-specific data for *e*, *p*, and *q*. We detrend the residuals for a given country by regressing them on a constant and a time trend; the resulting residuals are the estimated *e*-shocks and *p*-shocks.

[22] Note that Ireland pegged its currency to the British pound until it joined the EMS in 1979. Ireland's exchange rate volatility increased in the mid-1980s as part of a stabilization program involving depreciation of the Irish pound.

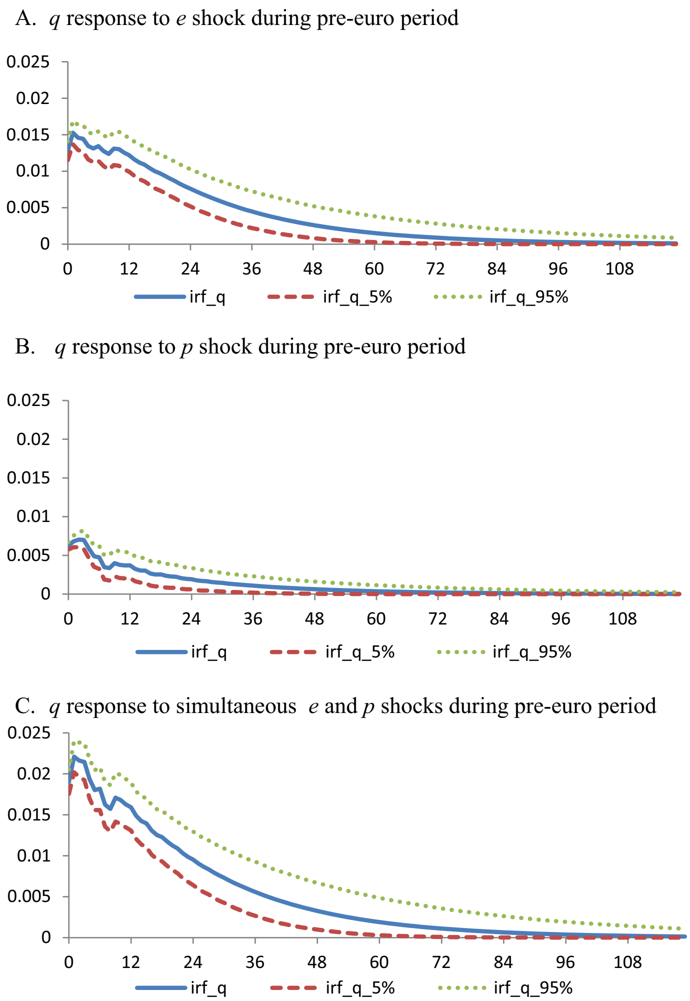

Fig. 3. The impulse response function (IRF) of the real exchange rate q in months conditional on one standard-deviation shocks of the exchange rate, prices, and both variables simultaneously, respectively, during the pre-euro period. Dashed lines are 5% and 95% confidence intervals, constructed using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

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 5. 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 4. 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 5 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.[23]

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

[23] 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: $\Delta q_t = \psi_0 q_{t-1} + \psi_1 \Delta q_{t-2} + \cdots + \psi_9 \Delta q_{t-9} + \varepsilon_{qt}$. 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 $i = 0, \ldots, 9$. Simulation of the restricted VECM system implies an unconditional half-life for the real exchange rate of 2.36 years, 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.

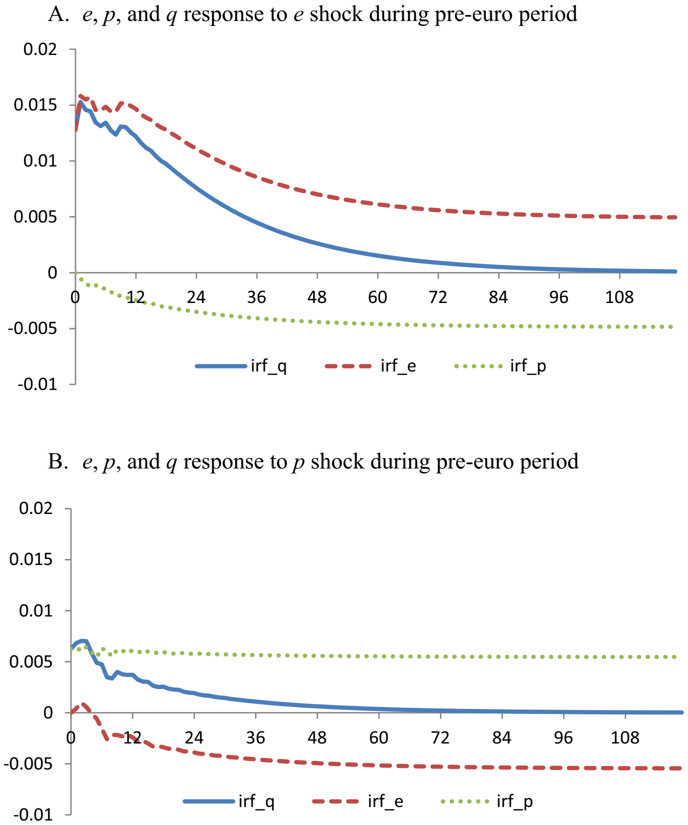

Fig. 4. The impulse response function (IRF) of the nominal exchange rate, the real exchange rate, and the price level in months conditional on one standard-deviation shocks of the nominal exchange rate and prices, respectively, during the pre-euro period based on bias-corrected CCEP estimates of the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

coefficients in the Δp equation to their values from Eq. (5) estimated during the euro period. The simulation makes use of the AR(10) estimates of Eq. (3.2) for p reported in Table 2 for the euro period to recover estimates of the γ_i coefficients in Eq. (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 $m = 1,2,\ldots,M_2 - 1$. Simulation 2 generates a half-life conditional on p shocks of 1.690 years (see column 2 of Table 5), very close to the value of 1.689 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 5 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 (37%) by which we found in Table 1 that 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.[24] 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 Fig. 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, Fig. 4 plots the IRF of the nominal exchange rate, real

[24] 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.

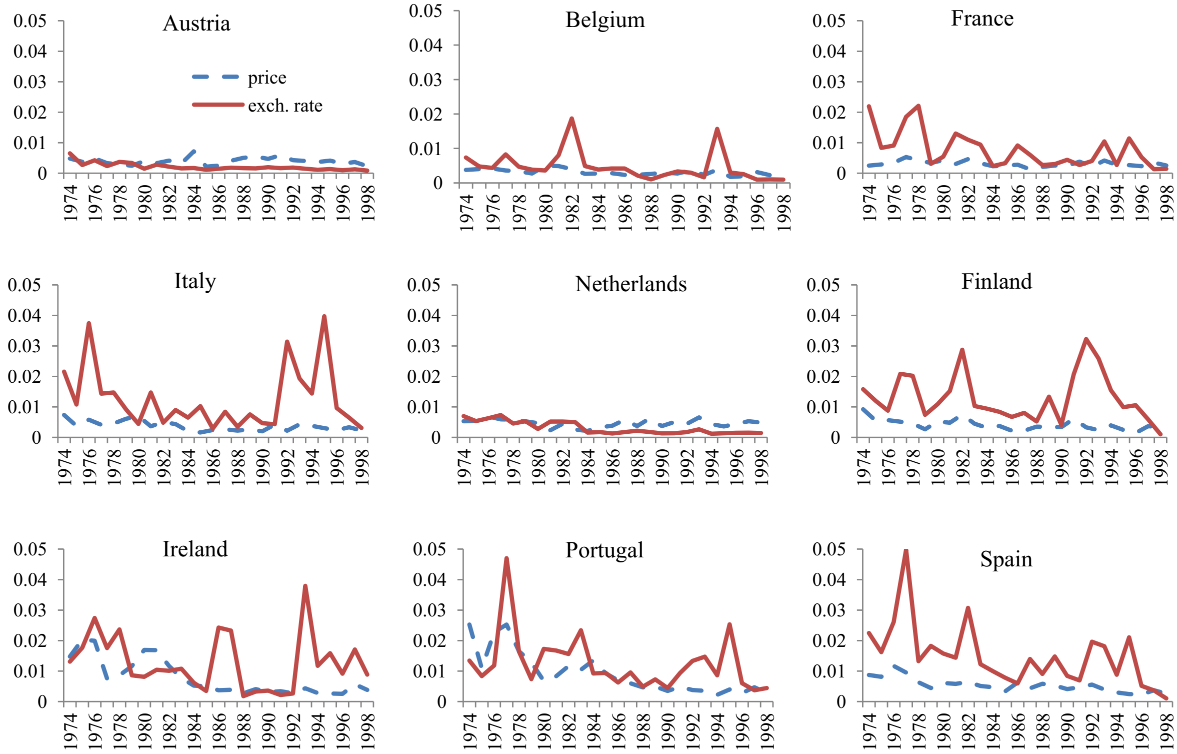

Fig. 5. Standard deviations of nominal exchange rate and price ratio shocks by country, constructed from the estimated VECM specification in Eq. (4). The nominal exchange rate shock is represented by the solid line; price shock by the dashed line.

Table 5
Counterfactual simulations.

	Simulation	Half-life of real exchange rate (in years)		
		e shock	p shock	Simultaneous e, p shocks
1.	Benchmark	2.331	1.139	2.002
2.	Nest AR of p estimated for the euro period: $\rho_{10} = a_{11} = \cdots = a_{1k-1} = b_{11} = \cdots = b_{1M-1} = 0, c_{11} = \cdots = c_{1M-1} = 0, d_{11} = \hat{\gamma}_1, \ldots d_{1M-1} = \hat{\gamma}_{1M-1}, \rho_{20} = \hat{\gamma}_0$	1.516	1.690	1.584
3.	Remove short-run response to Δe_{t-m} in Δe_t equation: $a_{11} = \cdots = a_{1M-1} = 0$	1.713	1.636	1.692
4.	Remove long-run response to q_{t-1} in Δe_t equation: $\rho_{10} = 0$	10.641	5.368	9.171
5.	Strengthen long-run response to q_{t-1} in Δp_t equation: $\rho_{20} = \hat{\gamma}_0$	1.537	0.543	1.271
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: $\rho_{20} = \hat{\gamma}_0, \rho_{10} = 0$	2.923	1.285	2.460

Note: Table reports counterfactual simulations based on VECM estimates reported for Eq. (4) for the pre-euro period in Table 4. Simulations 2, 5, and 6 make use of non-nested AR estimates of Eq. (3.2) for p reported in Table 2 for the euro period in order to recover estimates of the γ_i coefficients in Eq. (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.

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: $a_{11} = a_{12} = \cdots = a_{1,M-1} = 0$. In this case, as reported in simulation 3 of Table 5, 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 5 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 5.37 years for price shocks and 2.33 to 10.64 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. Recall 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 Eq. (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 4 indicates that the estimated value of ρ_{20} during the pre-euro period (-0.01) reported in Table 4 is smaller in absolute value than that of γ_0 in Eq. (5) during the euro period (-0.038) implicitly estimated from the coefficients reported in Table 2. 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.[25] 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 $\hat{\gamma}_0$ (=-0.038), and find that the half-life conditional on both exchange rate and price shocks falls by 37% (from 2.00 in the benchmark to 1.27). 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 5 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 1.29, even below that estimated for the euro period from the autoregression in Table 2 (1.69). Thus, even though losing the exchange rate as an adjustment mechanism can dramatically amplify the half-life (10.64 in simulation 4 of Table 5), 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 employ a 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

[25] This hypothesis is also tested in Huang and Yang (2015). We find that the elimination of exchange rate shocks is just as important.

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.

Appendix A. Supplementary material

Supplementary data associated with this article can be found, in the online version, at http://dx.doi.org/10.1016/j.jimonfin.2017.02.005.

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