

Deviations from Purchasing Power Parity Under Different Monetary Regimes: Do They Revert and, If So, How?

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Abstract

A vast literature has reported that deviations from long-run purchasing power parity (PPP) either show no tendency to die out or dissipate at an implausibly slow speed, generating the so-called ‘PPP puzzle.’ These studies do not account for at least one of three relevant issues: (i) adjustment toward PPP may occur via nominal exchange rates and relative prices at different speeds; (ii) different monetary and exchange rate regimes may generate regime shifts or breaks in the structural dynamics of PPP deviations; (iii) both theoretical and empirical research suggest that PPP deviations display significant nonlinearities. In this paper we employ an empirical model that accounts for all of these issues, hence encompassing much previous empirical research. Using a century of data for the G7 countries, we provide strong evidence that long-run PPP is valid, the relative importance of nominal exchange rates and prices in restoring PPP varies over time and across different monetary regimes, and reversion to PPP occurs at a speed that is consistent with the nominal rigidities suggested by conventional open economy macroeconomic models.

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1 Introduction

1.1 Overview

The purchasing power parity (PPP) hypothesis states that national price levels should be equal when expressed in a common currency. Although very few economists would believe that this simple proposition holds at each point in time, a large literature in international finance has examined empirically the validity of PPP over the long-run either by testing whether nominal exchange rates and relative prices move together in the long run or by testing whether the real exchange rate has a tendency to revert to a stable equilibrium level over time. The latter approach is motivated by the fact that the real exchange rate may be defined as the nominal exchange rate adjusted for relative national price levels, and therefore variations in the real exchange rate represent deviations from PPP, which must be stationary if long-run PPP holds (see the surveys of Froot and Rogoff, 1995; Rogoff, 1996; Sarno and Taylor, 2002).

Although long-run PPP is such a simple proposition about exchange rate behavior, it is not surprising that it has attracted the attention of researchers for decades. Indeed, whether long-run PPP holds or whether the real exchange rate is stationary has important economic implications on a number of fronts. In particular, the degree of persistence in the real exchange rate can be used to infer what the principal impulses driving exchange rate movements are. For example, if the real exchange rate is highly persistent or close to a random walk, then the shocks are likely to be real-side, principally technology shocks, whereas if it is not very persistent, then the shocks must be principally to aggregate demand, such as, for example, innovations to monetary policy (Rogoff, 1996). Further, from a theoretical perspective, if PPP is not a valid long-run international parity condition this casts doubts on the predictions of much open economy macroeconomics that is based on the assumption of long-run PPP. Indeed, the implications of open economy dynamic models are very sensitive to the presence or absence of a unit root in the real exchange rate (e.g. Lane, 2001; Sarno, 2001). Finally, estimates of PPP exchange rates are often used for practical purposes such as determining the degree of misalignment of the nominal exchange rate and the appropriate policy response, the setting of exchange rate parities, and the international comparison of national income levels. These practical uses of the PPP concept, and in particular the calculation of PPP

exchange rates, would obviously be of very limited use if PPP deviations contain a unit root.

Regardless of the great interest in this area of research, manifested by the large number of papers on PPP published over the last few decades, and regardless of the increasing quality of data sets utilized and of the econometric techniques employed, the validity of long-run PPP and the properties of PPP deviations remain the subject of ongoing controversies. Specifically, earlier cointegration studies generally reported the absence of significant mean reversion of the real exchange rate for the recent floating experience (Taylor, 1988; Mark, 1990), but were supportive of reversion toward PPP for the gold standard period (McCloskey and Zecher, 1984; Diebold, Husted and Rush, 1991), for the interwar float (Taylor and McMahon, 1988), for the 1950s US-Canadian float (McNown and Wallace, 1989), and for the exchange rates of high-inflation countries (Choudhry, McNown and Wallace, 1991). More recent applied work on long-run PPP among the major industrialized economies has, however, been more favorable toward the long-run PPP hypothesis for the recent float (e.g. Corbae and Ouliaris 1988; Cheung and Lai, 1993a, 1993b, 1994, 1998; Frankel and Rose, 1996).

One well-documented explanation for the inability to find evidence of a long-run relationship between the nominal exchange rate and PPP is the low power of conventional statistical tests to reject a false null hypothesis of a unit root in the real exchange rate or no cointegration between the nominal exchange rate and relative prices with a sample span corresponding to the length of the recent float (Frankel, 1986, 1990; Froot and Rogoff, 1995; Lothian and Taylor, 1997). Researchers have sought to overcome the power problem in testing for mean reversion in the real exchange rate either through long span studies (e.g. Kim, 1990; Lothian and Taylor, 1996; Taylor, 2002) or through panel unit root studies (e.g. Abuaf and Jorion, 1990; Frankel and Rose, 1996; O'Connell, 1998; Papell, 1998; Sarno and Taylor, 1998; Taylor and Sarno, 1998) or through time-series models that account for the possibility of nonlinear mean reversion toward PPP (e.g. Michael, Nobay and Peel, 1997; Obstfeld and Taylor, 1997; Taylor, 2001; Taylor, Peel and Sarno, 2001; Nakagawa, 2002). However, whether or not the long-span or panel-data studies do in fact answer the question whether PPP holds in the long run remains contentious (e.g. Engel, 1999, 2000). As far as the long-span studies are concerned, as noted in particular by Frankel and Rose (1996), the long samples required to generate a reasonable level of statistical power with standard univariate unit root tests

may be unavailable for many currencies (perhaps thereby generating a ‘survivorship bias’ in tests on the available data - Froot and Rogoff, 1995) and, in any case, may potentially be inappropriate because of differences in real exchange rate behavior both across different historical periods and across different nominal exchange rate regimes (e.g. Baxter and Stockman, 1989; Hegwood and Papell, 1999; Taylor, 2002). As for panel-data studies, these provide mixed evidence. While, for example, Abuaf and Jorion (1990), Frankel and Rose (1996) and Taylor and Sarno (1998) find results favorable to long-run PPP, O’Connell (1998) and Papell (1998) reject it on the basis of their empirical evidence. With respect to studies using nonlinear models of the real exchange rate, they unanimously support the validity of long-run PPP (see the review of Sarno and Taylor, 2002, and the references therein). However, since they are based on univariate regressions for the real exchange rate and do not allow for the possibility that the dynamics of PPP deviations be affected by monetary regimes, to date these studies have not shed light on whether nominal exchange rates or prices drive the adjustment toward the PPP equilibrium and do not investigate the role of different monetary and exchange rate regimes on the behavior of PPP deviations.

1.2 Questions addressed and approach

In light of the evidence provided by this literature, there remain at least three important issues in this area of research. First, it is still controversial whether long-run PPP is validated by the data. Second, it is debatable whether, when PPP is validated by the data, adjustment toward the long-run equilibrium level defined by this relationship is driven primarily by the exchange rate, by relative prices, or by both of them. Third, it is puzzling why the majority of studies find empirical estimates of the persistence of PPP deviations that are too high to be explained in light of conventional nominal rigidities and cannot be reconciled with the large short-term volatility of real exchange rates.

Our empirical analysis is devoted to shed light on all of these three issues. We start from noting three features that we view as potentially important in designing a suitable model for the deviations from PPP. The first feature is that the model needs to allow for the fact that adjustment toward PPP is likely to occur at different speeds via nominal exchange rates and prices. The vast majority of empirical studies on PPP is based on univariate representations of the real exchange

rate.¹ This approach is only valid if certain common factor restrictions in the unknown data generating process linking exchange rates and prices are satisfied. We find that these common factor restrictions are generally strongly rejected by the data, implying a loss of power in testing the null hypothesis that the real exchange rate is nonstationary (or that PPP is invalid) in conventional testing procedures. Employing a model which does not impose these restrictions increases test power, whilst allowing us to shed light on the relative importance of nominal exchange rates and prices in restoring the PPP equilibrium. The second desirable feature is that the model allows explicitly for the possibility that different monetary and exchange rate regimes may generate regime shifts or breaks in the structural dynamics of PPP deviations, especially when using long spans of data. The third feature is that the model might be nonlinear, in accordance with the growing evidence that exchange rate dynamics displays statistically and economically important regime shifts and nonlinearities.² Indeed, a related literature, which we take seriously into account in this paper, has provided mounting evidence that the conditional distribution of nominal and real exchange rate changes is well described by a mixture of normal distributions and that a Markov-switching model may be a good characterization of exchange rate behavior (e.g. see Engel and Hamilton, 1990; LeBaron, 1992; Engel, 1994; Engel and Hakkio, 1996; and Engel and Kim, 1999; Dueker and Neely, 2002; Clarida, Sarno, Taylor and Valente, 2003).

In this paper we extend the long-span data used by Obstfeld and Taylor (2003) and apply a very general modelling methodology in which regime changes in the data generating process are explicitly allowed for, focusing on the G7 countries. Taylor (2002) has examined these time series in order to establish the validity of PPP across a broad range of countries in an historical perspective, sticking to single-equation and panel linear econometric methods. Among his conclusions, which are supportive of long-run PPP, there is the notion that PPP deviations have similar half-lives across the different monetary regimes of the last century, but much larger shocks to the real exchange rate process have characterized floating exchange rate regimes relative to fixed exchange rate regimes. Our paper builds on and extends the work of Taylor (2002) and other scholars who

¹Exceptions which have considered this issue include, *inter alia*, Edison (1987), Wei and Parsley (1995), Edison, Gagnon and Melick (1997), Goldfajn and Valdes (1999), Coakley and Fuertes (2000), Engel and Morley (2002).

²For a discussion of the type of nonlinear models proposed by literature in this context, see the relevant section in the survey of Sarno and Taylor (2002) and the references therein.

have contributed to the literature on long-span real exchange rate behavior - including Edison (1987), Kim (1990), Lothian and Taylor (1996), Michael, Nobay and Peel (1997) - in several directions. We re-examine whether long-run PPP is valid and measure the relative importance of the exchange rate and the relative price in driving the adjustment toward long-run equilibrium. In particular, we investigate the relative importance of exchange rates and relative prices in restoring PPP across different exchange rate regimes, including the gold standard, the Bretton-Woods period, and the recent float. With over a century of data and different exchange rate arrangements over our sample period, we tentatively hypothesize that during floating exchange rate periods, the nominal exchange rate may be relatively more important in restoring departures from long-run equilibrium, while the relative price should restore long-run equilibrium during fixed exchange rate periods.³ Over the sample period examined, the economic history of the countries involved has seen a number of fundamental changes in monetary and exchange rate regimes, institutional structure and policy targets which, in addition to the continuous evolution of the financial system and various nominal and real shocks, represent serious potential pitfalls to researchers attempting to find an empirical model of the deviations from PPP that is stable over the full sample. The time-invariant, linear framework generally adopted by the literature may not be suitable for the purposes of this paper. We investigate whether allowing for regime-switching in the underlying data-generating process is an adequate characterization that is capable of capturing the impact of the different monetary regimes of the last century on the dynamics of exchange rates and relative prices. This is done through estimating a fairly general Markov-switching vector error correction model (MS-VECM) which characterizes the dynamic relationship between exchange rates and relative prices allowing for regime shifts in intercept, the variance-covariance matrix, and the entire set of parameters (autoregressive components and the cointegrating matrix).

³It is important to note that the latter statement is subject to the caveat that during fixed exchange rate regimes devaluations may be used by policy makers to induce adjustments in the real exchange rate. Conversely, during a floating regime, policy makers may use heavily foreign exchange market intervention for the purpose of reducing the variability of the real exchange rate (e.g. Sarno and Taylor, 2001; Calvo and Reinhart, 2002). Given these caveats, it may be that, in fact, irregular shifts in nominal exchange rates may be responsible for inducing reversion to PPP in fixed exchange rate regimes or, vice versa, the role of nominal exchange rates in driving the adjustment to PPP during floating regimes is not more important than the role of relative prices. Hence, the question of whether and how PPP deviations dissipate remains an empirical issue.

1.3 Main results

The results are supportive of long-run PPP for each exchange rate examined and of our basic conjecture: we find that during fixed exchange rate regimes, relative prices adjust to restore deviations from long-run equilibrium, while exchange rates bear most of the burden of adjustment during flexible exchange rate regimes. The estimated transition probabilities are consistent with the general result that the relative importance of exchange rates and relative prices in restoring the long-run equilibrium level of the exchange rate varies over time and is affected by the nominal exchange rate arrangement in operation.

Further, the estimated half-lives of the regime-dependent persistence profiles are sensibly different for fixed and floating regimes. During fixed exchange rate regimes the effects of shocks on the PPP equilibrium relationship may last for more than two years on average with a confidence interval comprising a minimum of 1.5 year and a maximum of almost 5 years. However, the corresponding half-lives during flexible exchange rate regimes are drastically shorter, since the nominal exchange rate is allowed to operate and contribute to restoring PPP. In fact, shocks will last for 0.8 years on average and the upper bound of the confidence interval is about 1 year.

We conclude that the PPP puzzles recorded in the literature may have been generated by the fact that none of the studies in the literature has explicitly allowed for all of the three features of PPP deviations described above, either imposing common factor restrictions to examine the stochastic properties of the real exchange rate rather than the dynamic interaction of nominal exchange rates and relative prices, or ignoring the impact of different monetary and exchange rate regimes on PPP deviations, or neglecting the possibility of significant nonlinearities. Allowing for all of these features simultaneously has allowed us to build a generalization of existing empirical frameworks that sheds new light on the behavior of PPP deviations. The properties of PPP deviations implied by our model appear to be consistent with standard models of open economy macroeconomics and with their dynamic properties under conventional nominal rigidities (e.g. Chari, Kehoe and McGrattan, 2002).

1.4 Organization

The remainder of the paper is set as follows. Section 2 defines long-run PPP and describes the dynamic relationship between exchange rates and relative prices, outlining the importance of common factor restrictions in this context. In Section 3 we set out the econometrics of Markov-switching multivariate models as applied to nonstationary processes and cointegrated systems. In Section 4 we describe our data set, while in the following section we report and discuss our empirical results, including unit root and cointegration tests, tests of common factor restrictions, linear and nonlinear VECM estimations, and the calculation of half-lives of shocks to PPP deviations using regime-dependent persistence profiles. A final section concludes.

2 Long-run purchasing power parity and the dynamic interaction between exchange rates and prices

Defining s_t as the logarithm of the nominal exchange rate (expressed as domestic price of foreign currency) and π_t as the logarithm of the ratio of domestic to foreign prices, then

$$q_t = s_t - \pi_t \tag{1}$$

may be seen as the deviation from PPP or the logarithm of the real exchange rate. Absolute long-run PPP would allow $q_t \neq 0$ in the short run, but it would require $q_t = 0$ in the long run. A less strict version of long-run PPP postulates that q_t may have a non-zero mean but it has to be a realization of a stationary process. If both the nominal exchange rate s_t and the relative price π_t have a stationary, invertible, non-deterministic ARMA representation after differencing once (i.e. $s_t, \pi_t \sim I(1)$), this definition of long-run PPP implies that s_t and π_t move together in the long run and exhibit a common stochastic trend, cointegrating with one cointegrating vector $[1, -1]$.

The relationship between s_t and π_t can be described by a cointegrating vector autoregression (VAR) of the form:

$$g(L)y_t = \varepsilon_t \tag{2}$$

where y_t is a two-dimensional observed time series vector, $y_t = [s_t, \pi_t]'$; $g(L)$ is a suitable p -th order, 2×2 matrix polynomial in the lag operator L ; $\varepsilon_t = [\varepsilon_{1t}, \varepsilon_{2t}]'$ is a vector of Gaussian white noise processes with covariance matrix Σ , $\varepsilon_t \sim NID(\mathbf{0}, \Sigma)$.⁴ By the Granger Representation Theorem (Engle and Granger, 1987), s_t and π_t must possess a vector error correction model (VECM) representation where the deviation from PPP, q_t (the real exchange rate) plays the part of the equilibrium error:

$$\Delta y_t = \Gamma(L)\Delta y_{t-1} + \Pi y_{t-1} + \varepsilon_t, \quad (3)$$

where $\Gamma(L)$ is a 2×2 matrix polynomial; and the long-run impact matrix $\Pi = \alpha\beta'$, where α and β are 2×1 vectors, with β denoting the cointegrating vector (assumed to be $[1, -1]$ under PPP) and α the vector of weights on the cointegrating vector in each of the two equations of the VECM. Note that $g(L)$ in equation (2) is unrestricted; hence, $\Gamma(L)$ and Π are also unrestricted. The existence of one cointegrating relationship between s_t and π_t implies that the rank of Π equals unity.

To understand the relationship between tests of PPP using the VECM (3) and unit root tests based on the augmented Dickey-Fuller (ADF) auxiliary regression or variants of it often applied by researchers to the real exchange rate, let us start from noting that the latter focuses on the roots of $q_t = \beta' y_t$ rather than the properties of y_t itself. This implicitly imposes certain common factor restrictions, as it can be seen by pre-multiplying (3) by β' to obtain

$$\beta' \Delta y_t = \beta' \Gamma(L) \Delta y_{t-1} + (\beta' \alpha) \beta' y_{t-1} + \beta' \varepsilon_t \quad (4)$$

or

$$[1 - G(L)L] \Delta q_t = \rho q_{t-1} + \psi_t \quad (5)$$

where the coefficient $\rho = \beta' \alpha$; $G(L)$ is a scalar polynomial in L , and

$$\psi_t = [\beta' \Gamma(L) - G(L)\beta'] \Delta y_{t-1} + \beta' \varepsilon_t. \quad (6)$$

⁴For ease of exposition, in this section we don't allow for a constant term in the cointegrating VAR and do not allow for regime-switching in the VAR. However, it is important to note that none of the points made below is dependent on this simplification. In fact, allowance for a constant term or generalization to a nonlinear VAR would simply increase the number of common factor restrictions on which we focus in this section.

Equation (5) is the conventional ADF regression used by a number of researchers for testing the null hypothesis of a unit root in the real exchange rate. It is now apparent that the disturbance term ψ_t may contain valuable information for two reasons.⁵ First, unless $\beta'\Gamma(L) = G(L)\beta'$, lags of Δy_t enter ψ_t . Second, if π_t is not weakly exogenous, then $\beta'\varepsilon_t$ may be explained partly by the current value of π . Both of these reasons imply a loss of information from testing for PPP by testing for a unit root in the real exchange rate q_t using an ADF regression of the form (5) rather than by analyzing the full VECM linking the nominal exchange rate s_t and the relative price π_t (3). In turn, as demonstrated for several different data generating processes by Kremers, Ericsson and Dolado (1992), this loss of information leads to a substantial loss of power of the ADF test in rejecting the null hypothesis of a unit root, which, in this context, would bias the outcome of the test toward concluding that the real exchange rate is nonstationary and long-run PPP is invalid.

The discussion in this section suggests yet another reason, unexplored by the literature to date, why unit root tests have low power in the context of testing for PPP. Several researchers have noted that conventional unit root tests have low power in rejecting the null of real exchange rate nonstationarity for the recent floating exchange rate regime since 1973 or so, because the sample size is too short to yield sufficient test power (see Froot and Rogoff, 1995). This has led several researchers to develop panel unit root tests which exploit the cross-correlation in exchange rate data in an attempt to increase test power (since at least Hakkio, 1984; and Abuaf and Jorion, 1990). However, these tests also impose the same common factor restrictions as single-equation unit root tests; in fact, the number of restrictions increases with the size of the panel. Indeed, if the common factor restrictions are not satisfied, for any given sample size, the power of (single-equation and panel) unit root tests may be substantially lower than the power of tests based on the full VECM because unit root tests applied to the real exchange rate effectively ignore valuable information by assuming ‘error’ rather than ‘structural’ dynamics.

We investigate the empirical relevance of these issues by using the VECM representation of s_t and π_t to test for PPP and to shed light on the relative importance of the nominal exchange rate and relative prices in restoring the PPP equilibrium level across different nominal regimes since

⁵This issue is a specific case of the a more general class of problems analyzed by several authors. Notably, see Kremers, Ericsson and Dolado (1992).

the 19th century. Further, in order to take seriously into account the possibility of regime shifts over long spans of data, we use a generalization of the standard linear VECM (3) which is capable of allowing all of the VECM parameters to change over time and to identify the various regimes that characterize the long sample periods examined.

3 Markov-switching error correction

In this section we outline the econometric procedure employed in order to model regime shifts in the dynamic relationship between spot exchange rates and relative prices. The procedure essentially extends Hamilton's (1988, 1989) Markov-switching regime framework to nonstationary systems, allowing us to apply it to cointegrated VAR and VECM systems (see Krolzig, 1997, 1999).⁶

Consider the following M -regime p -th order Markov-switching vector autoregression (MS(M)-VAR(p)) which allows for regime shifts in the intercept term:

$$g(L)y_t = \nu(z_t) + \varepsilon_t, \quad (7)$$

where y_t is a K -dimensional observed time series vector, $y_t = [y_{1t}, y_{2t}, \dots, y_{Kt}]'$; $\nu(z_t)$ is a K -dimensional column vector of regime-dependent intercept terms, $\nu(z_t) = [\nu_1(z_t), \nu_2(z_t), \dots, \nu_K(z_t)]'$; $g(L)$ is a suitable p -th order, $K \times K$ matrix polynomial in the lag operator L ; $\varepsilon_t = [\varepsilon_{1t}, \varepsilon_{2t}, \dots, \varepsilon_{Kt}]'$ is a K -dimensional vector of Gaussian white noise processes with covariance matrix Σ , $\varepsilon_t \sim NID(\mathbf{0}, \Sigma)$. The regime-generating process is assumed to be an ergodic Markov chain with a finite number of states $z_t \in \{1, \dots, M\}$ governed by the transition probabilities $p_{ij} = \Pr(z_{t+1} = j \mid z_t = i)$, and $\sum_{j=1}^M p_{ij} = 1 \forall i, j \in \{1, \dots, M\}$.

A standard case in economics and finance is that y_t is nonstationary but first-difference stationary, i.e. $y_t \sim I(1)$. Then, given $y_t \sim I(1)$, there may be up to $K - 1$ linearly independent cointegrating relationships, which represent the long-run equilibrium of the system, and the equilibrium error (the deviation from the long-run equilibrium) is measured by the stationary stochastic

⁶The class of nonlinear models which could be used to characterize a nonlinear error correction model is infinite, and we have chosen to concentrate on the Markov-switching VECM primarily because of its generality, its relative simplicity, and the large amount of previous research on the estimation of MS-VECMs. An alternative approach might be to consider, for example, a threshold VECM or a smooth transition VECM, which would generalize the threshold autoregressive and the smooth transition autoregressive models used by several researchers in this context.

process $u_t = \alpha' y_t - \beta$ (Granger, 1986; Engle and Granger, 1987). If indeed there is cointegration, the cointegrated MS-VAR (7) implies a Markov-switching vector error correction model or MS-VECM of the form:

$$\Delta y_t = \nu(z_t) + \Gamma(L)\Delta y_{t-1} + \Pi y_{t-1} + \varepsilon_t, \quad (8)$$

where $\Gamma(L)$ is a suitable $K \times K$ matrix polynomial in the lag operator L of order $p - 1$, such that $\Gamma_i = -\sum_{j=i+1}^p g_j$ for $i = 1, \dots, p - 1$ are matrices of parameters, and $\Pi = \sum_{i=1}^p \Pi_i - \mathbf{I}$ is the long-run impact matrix whose rank r determines the number of cointegrating vectors (e.g. Johansen, 1995; Krolzig, 1999).

Although, for expositional purposes, we have outlined the MS-VECM framework for the case of regime shifts in the intercept alone, shifts may be allowed for elsewhere. The present application focuses on a multivariate model comprising, for each of the countries analyzed, the spot exchange rate and the relative price (hence $y_t = [s_t, \pi_t]'$), for which, following the reasoning of Section 2, one unique, independent cointegrating relationship, represented by the deviation from PPP q_t , should exist. As discussed in Section 5 below, in our empirical work, after considerable experimentation, we selected a specification of the MS-VECM which allows for regime shifts in the intercept, the variance-covariance matrix and the whole set of parameters (autoregressive component, $\Gamma(L)$ and cointegration matrix, Π). This model, the Markov-Switching-Intercept-Autoregressive-Heteroskedastic-VECM or MSIAH-VECM, may be written as follows:

$$\Delta y_t = v(z_t) + \Gamma(L)(z_t)\Delta y_{t-1} + \Pi(z_t)y_{t-1} + u_t, \quad (9)$$

where $\Pi(z_t) = \alpha(z_t)\beta'$, $u_t \sim NIID(\mathbf{0}, \Sigma(z_t))$ and $z_t \in \{1, \dots, M\}$.

An MS-VECM can be estimated using a two-stage maximum likelihood procedure. The first stage of this procedure essentially consists of the implementation of the Johansen (1988, 1991) maximum likelihood cointegration procedure in order to test for the number of cointegrating relationships in the system and to estimate the cointegration matrix. In fact, in the first stage use of the conventional Johansen procedure is legitimate without modelling the Markovian regime shifts explicitly (see Saikkonen, 1992; Saikkonen and Luukkonen, 1997). The second stage then consists of the implementation of an expectation-maximization (EM) algorithm for maximum likelihood

estimation which yields estimates of the remaining parameters of the model (Dempster, Laird and Rubin, 1977; Hamilton, 1993; Kim and Nelson, 1999; Krolzig, 1999).

We now turn to a brief discussion of our data set and then to our empirical analysis.

4 Data

The data set used in this study comprises annual observations for the nominal exchange rate (domestic price of foreign currency) and the price levels - based on the consumer price index (CPI) or the gross domestic product (GDP) deflator, depending on availability - relative to the US for each of the G7 countries. Our data set is obtained from updating the relevant time series in the data set constructed by Obstfeld and Taylor (2003) using the *International Financial Statistics* of the International Monetary Fund.⁷ From these data we calculate the logarithm of the nominal exchange rate, s_t , the relative price, π_t , and the real exchange rate, $q_t = s_t - \pi_t$, as defined in equation (1).

The sample spans from the late 19th or early 20th century to the late 20th century and thus covers a variety of international monetary arrangements, including the classical gold standard, the Bretton Woods era, the modern float and, for some countries, the Exchange Rate Mechanism of the European Monetary System. The exact sample period for each country is as follows: 1870-2000 for Canada and the UK; 1880-1998 for France, Germany and Italy; 1885-2000 for Japan. The start date was in each case dictated by data availability, whereas the end date is 2000 except for the three G7 countries which joined the European Monetary Union and replaced their respective national currencies with the euro in January 1999.

5 Empirical results

5.1 Unit roots, cointegration, and common factor restrictions

As a preliminary exercise, we test for unit root behavior of the nominal exchange rate (s_t), the relative price (π_t) and the real exchange rate (q_t) time series, for each of the six dollar exchange rates under investigation, employing two efficient unit-root tests (see Elliot, Rothemberg and Stock,

⁷We are very thankful to Alan Taylor for graciously providing the data. Our updated data set is available upon request.

1996; Cheung and Lai, 1998). In Table 1 we report the results from calculating the MZ_α and MZ_t tests proposed by Ng and Perron (2001); these tests use generalized least squares-detrending to maximize test power and a modified information criterion to select the lag truncation in order to minimize size distortion.

Consistent with the results reported by Taylor (2002, Table 3), we are unable to reject the unit root null hypothesis at conventional nominal levels of significance for the real exchange rates of Japan and Canada, but we can reject the null for the real exchange rates of Germany, Italy and the UK as well as, albeit marginally and only under the MZ_t test, for France. With respect to the nominal exchange rate and the relative price time series, in no case we can reject the null of a unit root. On the other hand, the first difference of each time series investigated is found to be clearly stationary.⁸ Hence, the unit root tests clearly indicate that each of the time series for the nominal exchange rate and relative price examined is a realization from a stochastic process integrated of order one. Also, we find some favorable evidence of real exchange rate stationarity for four out of six exchange rates, although the rejection of the unit root null hypothesis is marginal for France and only occurs under the MZ_t test. Overall, we view the evidence on the validity of long-run PPP as mixed or, at best, mildly favorable. Nevertheless, we are aware that unit root tests applied to the real exchange rate may have low power for a variety of reasons, including the imposition of possibly invalid common factor restrictions.⁹

We then employ the Johansen (1988, 1991) maximum likelihood procedure in a VAR for $y_t = [s_t, \pi_t]'$, which does not impose any common factor restrictions.¹⁰ Indeed, we tested the common factor restrictions required for the VAR to reduce to a standard univariate autoregression for the real exchange rate, using likelihood ratio (LR) tests. The results, reported in Table 2, indicate that, in each case, the restrictions are strongly rejected, providing the case for using unrestricted VAR and VECM estimation for testing PPP rather than real exchange rate autoregressions.

⁸The results for differenced series are not reported to conserve space but are available upon request.

⁹Taylor (2002, Table 3) applies similar unit root tests on other real exchange rate series and provides evidence of mean reversion for many of them. In this paper we focus only on a specific subset of major developed economies, namely the G7 countries. For these countries, our Table 1 qualitatively replicates Taylor's findings.

¹⁰We allowed for a maximum lag length of four and chose, for each dollar exchange rate, the appropriate lag length on the basis of conventional information criteria. In each VAR we allowed for an unrestricted constant term, and in the VARs for Canada, France, Germany and Japan we also allow for a linear trend restricted to the cointegration space. However, the results we report were found to be qualitatively identical when using different specifications of the deterministic components in the VAR tested for cointegration.

On the basis of the Johansen likelihood ratio test statistics for cointegrating rank (based on the maximal eigenvalue and on the trace of the stochastic matrix) reported in Table 3, we again find mixed evidence for long-run PPP. In fact, only for three out of six exchange rates examined - namely Canada, France and Italy - we can reject the null hypothesis of no cointegration at the five percent significance level. Also, for the three exchange rates for which we find evidence in favor of PPP, when testing for the over-identifying restrictions on the β' matrix of cointegrating coefficients suggested by the framework discussed in Section 2, we reject the null hypothesis that $\beta' = [1 \ -1]$ in each case except for Italy.

There are a variety of reasons for the failure to find a unique cointegrating relationship between economic time series where one would normally be expected on the basis of economic theory (see Siklos and Granger, 1997, and the references therein). In particular, it is well known that the presence of structural breaks or regime shifts may perversely affect cointegration tests, which appears to be especially likely when using long spans of data. Such shifts may significantly alter the dynamic relationship that may exist between the variables, and tests of the long term behavior of these variables should account for them.

Hence, as a check of adequateness of the models as well as an additional motivation for employing a more general modelling framework, we test for the stability of the cointegrating rank over time. In particular, we carry out two recently developed tests that are specifically designed to test for stability within a cointegrating framework: the Hansen and Johansen (1999) test of the null hypothesis of stability of the eigenvalues of the stochastic matrix¹¹; and the test statistic developed by Quintos (1997) for testing the null hypothesis that the cointegrating rank is constant among breakpoints.¹² The test results, reported in the last two columns of Table 3, suggest that the null

¹¹The test carried out is the ‘fluctuation’ test of the eigenvalues of the stochastic matrix (Hansen and Johansen, 1999, p. 319). The null hypothesis being tested is that the cointegrating rank is stable over time. As demonstrated by Hansen and Johansen (1999), the asymptotic distribution of this test is nonstandard and is a function of the dimension of the cointegrating VAR, the number of cointegrating vectors, and the specified deterministic components.

¹²This procedure allows us to test whether the number of cointegrating vectors varies across sub-sample periods. If the rank varies, then the number of long-term relationships changes across break points, and Quintos’s procedure allows us to test a wide variety of null hypotheses. The null being tested in this paper allows for just one breakpoint, $J = 1$, endogenously determined as in Quintos (1995), and $H_0 : r_1 = r_2 = r$ where r_1 and r_2 are the cointegrating ranks in the pre- and post-breakpoint periods respectively. In the paper $H_0 : r_1 = r_2 = 1$. The null is tested by means of likelihood ratio (LR) statistics whose distributions are functions of scaled, K -dimensional Brownian motions and depend upon the number of variables in the cointegrating VAR, K , and the cointegrating ranks in the

hypothesis of temporal stability of the cointegrating rank is strongly rejected, further motivating the use of a regime-switching framework, to which we now turn.

5.2 Regime shifts and model selection

We next estimate a standard linear VECM using full-information maximum likelihood (FIML) methods:

$$\Delta y_t = \nu + \Gamma(L)\Delta y_{t-1} + \Pi y_{t-1} + u_t \quad (10)$$

where $y_t = [s_t, \pi_t]'$, employing a maximum lag of $p = 3$, chosen on the basis of the Akaike Information Criterion (AIC) and the Schwartz Information Criterion (SIC). Employing the conventional general-to-specific procedure, we obtained fairly parsimonious models for each exchange rate with no significant residual serial correlation.¹³ We then proceed to investigate the presence of nonlinearities and regime shifts further through the estimation of a fairly general Markov-switching model of the form:

$$\Delta y_t - \delta(z_t) = \alpha(z_t) [\beta' y_{t-1} - \mu(z_t)] + \Gamma(L)(z_t) [\Delta y_{t-i} - \delta(z_t)] + \omega_t, \quad (11)$$

where $y_t = [s_t, \pi_t]'$, $\delta(z_t)$ is the (2×1) regime-dependent vector of means of the short-run dynamics, and $\mu(z_t)$ is the regime-dependent mean of the long-run equilibrium relationships.

Next we apply the conventional ‘bottom-up’ procedure designed to detect Markovian shifts in order to select the most adequate characterization of an M -regime p -th order MS-VECM for Δy_t .¹⁴ The VARMA representations of the time series (Poskitt and Chung, 1996) suggests in each case that the number of regimes is in the range between two and four.

However, for any MS-VECM estimated the implicit assumption that the regime shifts affect alternatively only the intercept term, the variance-covariance matrix or the autoregressive compo- pre- and post-breakpoint periods r_1 and r_2 .

¹³ Full details on these estimation results are available from the authors upon request.

¹⁴ Essentially, the bottom-up procedure consists of starting with a simple but statistically reliable Markov-switching model by restricting the effects of regime shifts on a limited number of parameters and checking the model against alternatives. In such a procedure, most of the structure contained in the data is not attributed to regime shifts, but explained by observable variables, consistent with the general-to-specific approach to econometric modelling. For a comprehensive discussion of the bottom-up procedure, see Krolzig (1997).

ment of the VECM was found to be inappropriate. In fact, we checked the relevance of all these components by using likelihood ratio tests of the type suggested by Krolzig (1997, p. 135-6). The results reported in the first three rows of Table 4 indicate very strong rejection of the null of no regime dependence, clearly suggesting that an MS-VECM that allows for shifts in the intercept, the variance-covariance matrix *and* the autoregressive component, namely an MSIAH-VECM(p), is the most appropriate model within its class in the present application.

Further, in the same spirit of the tests previously computed, we carry out a further LR test in order to select the most parsimonious VECM for characterizing the dynamic relationship between spot exchange rates and relative prices. In particular, we test the null of MSIH-VECM(1) against the alternative of MSIH-VECM(p). This additional test is applied only to Italy since for the other countries a lag length of unity was already carried over from the linear estimation. As it may be seen in the fourth row of Table 4 we are not able to reject this null hypothesis at standard significance levels, hence concluding that one lag is appropriate for each MS-VECM estimated.

In order to discriminate between models allowing for two regimes against models governed by a higher number of regimes we also execute a further likelihood ratio test. It is well known that the presence of nuisance parameters gives the likelihood surface sufficient freedom so that the finding of some statistically significant parameters could simply be due to sampling variation. The scores associated with parameters of interest under the alternative hypothesis may be identically equal to zero under the null. In order to avoid this inconvenience several testing procedures have been proposed.¹⁵ The results reported in Table 4 (fifth and sixth rows) show very large test statistics and the corresponding p -values, calculated as in Ang and Bekaert (1998), suggest that three regimes may be appropriate in all cases, with the exception of France and Japan where four regimes are required to describe the dynamics of nominal exchange rates and relative prices. Finally, the

¹⁵It is important to note here that the regularity conditions under which the Davies (1977, 1987) test is valid are violated, since the Markov model has both a problem of nuisance parameters and a problem of ‘zero score’ under the null hypothesis. Moreover, even if the Davies bound suggested by Krolzig is appropriate, it is possible that it will only be valid if the null model is a linear model with iid errors; in the present case, it is difficult to believe that this condition is met since exchange rate innovations are not homoskedastic, which would induce some distortion. Therefore, the distribution of the LR test is likely to differ from the adjusted χ^2 distribution proposed by Davies (1977, 1987), and this is why we do not report marginal significance levels for the LR tests. For extensive discussions of the problems related to LR testing in this context, see Hansen (1992, 1996) and Garcia (1998). See Garcia and Perron (1996) for an empirical application.

linearity tests, reported in the last column of Table 4, indicate in each case the rejection of the linear VECM(p) in favor of its nonlinear, Markov-switching counterpart.

Hence, the final result of this procedure identifies for all countries an MS-VECM governed by three or four different regimes that can be written as follows:

$$\Delta y_t = v(z_t) + \Pi(z_t) y_{t-1} + \Gamma(L)(z_t) \Delta y_{t-1} + \omega_t, \quad (12)$$

where $\Pi(z_t) = \alpha(z_t) \beta'$, $\omega_t \sim NIID(\mathbf{0}, \Sigma(z_t))$ and $z_t = 1, 2, 3, 4$. We estimate the MSIAH-VECM (12) using an EM algorithm for maximum likelihood (Dempster, Laird and Rubin, 1977), for each of the countries under investigation.

The estimation yields fairly plausible estimates of the coefficients for the VECMs estimated, including the regime-dependent adjustment coefficients in $\alpha(z_t)$, which are generally found to be statistically significantly different from zero in at least one regime.^{16 17} Before turning to the empirical results from estimating the MS-VECMs, however, we report tests of regime-dependent tests of the common factor restrictions which would make the MS-VECM (12) reduce to a regime-switching real exchange rate autoregression of the type used, for example, by Engel and Hamilton (1990). Again, for each regime and each exchange rate examined, the restrictions are generally rejected, strengthening the case for testing for PPP in the context of a full VECM rather than in the context of unit root tests of nonstationarity of the real exchange rate. The likelihood ratio

¹⁶We also looked at graphs of standardized residuals, the smoothed residuals and the one-step prediction errors from each estimated MSIAH-VECM. The difference is concerned with the weighting of the residuals. Loosely speaking, the smoothed residuals are the closest to the sample residuals from a linear regression model; however, they overestimate the explanatory power of the Markov-switching model due to the use of the full-sample information covered in the smoothed regime vector. The standardized residuals are conditional residuals. The one-step prediction errors are based on the predicted regime probabilities. Unfortunately, many conventional diagnostic tests, such as standard residual serial correlation tests, may not have their conventional asymptotic distribution when the residuals come from Markov-switching models and are therefore not reported here. However, the graphs of standardized residuals, the smoothed residuals and the one-step prediction errors provided no visual evidence of residual serial correlation in any of the residuals series plotted.

¹⁷Although the main focus of the paper is not the comparison between linear and nonlinear VECMs we calculated the goodness-of-fit measures for the MSIAH-VECM and its corresponding linear counterpart, computing the ratio of the $\bar{R}^2$, the residual variance (RV), the AIC and the SIC from each of the estimated MSIAH-VECM selected by the bottom-up procedure, as previously discussed and reported in Table 3. The results (see Table A2 in the appendix) show that, for each of the countries examined, the estimated MSIAH-VECM outperforms the best alternative linear model, leading to a substantial reduction of the residual variance - more than 60% for Germany - and a remarkable improvement of the $\bar{R}^2$, in all countries higher than 12%.

tests results, reported in Table 5, indicate rejections of the common factor restrictions in at least two regimes for each MS-VECM with fairly low p -values.

5.3 MS-VECM estimation results

Table 6 reports the average smoothed probabilities relative to the different estimated regimes over four sub-sample periods. In particular we calculated the average smoothed probabilities over the Gold Standard (1870-1914), interwar (1915-1944), Bretton Woods (1945-1972) and recent float (1973-2000) periods in order to see whether our preferred MSIAH-VECMs were able to correctly identify different nominal exchange rate regimes. The results in Table 6 display several interesting patterns. First, the estimated MSIAH-VECMs are able to disentangle different nominal regimes with quite high precision. In fact, the reported regime classification measure (RCM) proposed by Ang and Bekaert (2002), which can be seen as a measure of the variance of the estimated Markov chain, is in all cases very close to zero. Second, the MSIAH-VECMs are able to identify regimes with fixed exchange rates from regimes with flexible exchange rates. In fact, in all cases we can see that the regime with the highest probability under the Gold Standard period is the same exhibiting the highest probability under the Bretton Woods period. Third, the interwar period is, across countries, the only sub-period not clearly identified by the MSIAH-VECMs. In fact the average probabilities are very similar across different regimes and the regime classification measure is at least 10 times higher than in the other sub-periods. This is not surprising given the instability of the exchange rate arrangements that characterizes the interwar period in many G7 economies and the fairly low number of observations for this period relative to the full sample.

This interpretation of regimes also benefits from examining jointly the results of Table 6 with the ones in Table 7, where we report the regime-dependent error correction coefficients. The results in Table 7 suggest that during fixed exchange rate regimes the error correction coefficients in the exchange rate equations are, generally, not significantly different from zero at conventional significance statistical levels, while the error correction coefficients in the relative price equations are statistically significant and correctly signed. This implies that during fixed exchange rate regimes the relative price responds to deviations from PPP in a way to restore the long-run exchange rate

equilibrium.¹⁸

Vice versa during flexible exchange rate regimes we find that the error correction coefficients in the exchange rate equations are strongly statistically significant while the error correction coefficients in the relative price equations, albeit correctly signed, are not always statistically different from zero. This is consistent with the fact that during flexible exchange rate regimes the exchange rate is primarily responsible for restoring deviations from long-run PPP equilibrium.

The above results are quite interesting. It is rare in the use of Markov-switching models that one finds such strong evidence of being in one particular regime or the other. The results of the MS-VECM in terms of classification of nominal regimes provide clear evidence that adjustment to long-run equilibrium takes different forms in fixed versus floating exchange rate regimes.

This finding appears to be in contrast, *prima facie*, with the findings of Engel and Morley (2002), who use a state-space representation of the VECM linking exchange rates and relative prices for the recent float alone and find that nominal exchange rates respond less than relative prices to PPP disequilibria. Note, however, that Engel and Morley focus on the speed of adjustment of exchange rate and prices toward their equilibrium levels (i.e. in response to what they term the ‘exchange rate gap’ and the ‘price gap’ respectively), rather than toward the PPP equilibrium level. In this sense, therefore, their estimates of speeds of adjustment are not inconsistent with and not comparable to the ones we report here, which focus directly on the response of exchange rates and prices to PPP deviations. Our results are also in line with the evidence provided by Wei and Parsley (1995), Goldfajn and Valdes (1999) and Cheung, Lai and Bergman (1999), who argue that the nominal exchange rate is responsible for most of the adjustment toward PPP, rather than nominal prices. However, our empirical framework, by generalizing this line of previous research to a nonlinear regime-switching VECM system, allows us to pin down the speed of adjustment toward PPP for a long-span of data and across regimes.

¹⁸While two of the regimes clearly identify fixed and floating exchange rate regimes, the third and fourth regimes do not lend themselves to any easy economic interpretation. However, inspection of the transition probabilities suggests that these regimes are largely characterized by outliers in the data, which do not fit easily in any of the first two regimes. Indeed, estimation of a more parsimonious two-regime VECM yielded similar results and economic implications as the three- or four-regime VECMs, but in a less clear cut fashion. This is because the allowance for more than two regimes, which was suggested by the bottom-up procedure, is able to remove the outliers in the data and to identify more precisely the fixed and floating exchange rate regimes, ultimately yielding a statistically superior model.

5.4 PPP deviations are not as puzzling as you think

While the estimated MS-VECMs impart some idea on the degree of mean reversion as well as the underlying mechanism governing the adjustment toward long-run PPP, a sensible way to gain a full insight into the mean-reverting properties of the estimated nonlinear models is through dynamic stochastic simulation. In particular, an analysis of the impulse response functions will allow the half-life of shocks to PPP deviations to be gauged and these can be compared to those previously reported in the literature in order to see if explicitly considering exchange rate regimes helps to resolve the PPP puzzle of very slow real exchange rate adjustment.

Thus, we examined the dynamic adjustment in response to shocks by means of the ‘persistence profiles approach’ developed by Pesaran and Shin (1996). This approach focuses on the analysis of the effect of system-wide shocks on equilibrium relationships within a cointegration framework, hence being a particularly appealing tool for measuring the response to PPP shocks in our context. In order to take into account the findings in the previous section, we generalize the persistence profile approach to the Markov-switching framework in order to get an insight on the effect of system-wide shocks on equilibrium relationships conditional on different regimes.¹⁹

The results are reported in Table 8. The estimated half-lives of the regime-dependent persistence profiles are sensibly different between the two regimes. During the fixed exchange rate regimes, when the relative price is primarily responsible for bearing the burden of adjustment toward PPP in response to shocks, the effects of system-wide shocks on the equilibrium relationships will last for more than two years on average with a confidence interval comprising a minimum of 1.5 year and a maximum of almost 5 years. However, the half-lives of system-wide shocks on the equilibrium relationships during flexible exchange rate regimes are drastically shorter. In fact, shocks will last for 0.8 years on average, and the upper bound of the confidence interval is about 1 year, as one might expect given that under flexible exchange rate regimes both the nominal exchange rate and the relative price can contribute to restoring PPP in response to shocks.

Overall, the fact that consensus estimates for the rate at which PPP deviations dumps, using long-span data or panel unit root tests, generally range between three and five years may be due to several reasons. First, several previous studies using long-span data acknowledged the

¹⁹See Appendix 2 for further details.

fact that the results were obtained by blending fixed and floating rate data, but to the best of our knowledge, this is the first attempt to explicitly allow for the presence of regime shifts in investigating PPP with long-span data across countries. Second, much previous research has employed (single-equation or panel, linear or nonlinear) autoregressions for the real exchange rate rather than focusing on the full VECM linking nominal exchange rates and relative prices, hence biasing the test outcome against long-run PPP. Our finding of very short half-lives for the recent float is presumably the product of taking into consideration simultaneously three issues that are relevant in determining the power of tests for PPP: examining the full VECM for exchange rates and prices rather than using real exchange rate autoregressions; using a long-span of data rather than only data for the recent float, while allowing for regime shifts; modelling the dynamics of PPP deviations in a nonlinear fashion.

6 Conclusions

In this paper we have re-examined the behavior of deviations from PPP using a century of data for the six dollar exchange rates obtaining among the G7 countries. We begun from noting the empirical success in establishing the validity of long-run PPP recorded by both long-span studies and studies based on nonlinear real exchange rate models, which contrasts with the unfavorable evidence provided by studies based on standard unit root tests and most cointegration tests as well as with the mixed evidence provided by studies employing panel unit root tests. However, long-span studies have often been accused of ignoring the possible regime shifts that may characterize PPP deviations across different monetary and exchange rate regimes. We also illustrated how linear and nonlinear real exchange rate autoregressions impose potentially invalid restrictions on the unknown data generating process driving the relationship between nominal exchange rates and prices, and how these restrictions may prevent us from detecting mean reversion toward PPP.

This process of interpretation of the time series evidence of the relevant literature led us to employ a nonlinear empirical framework for PPP deviations which allows us to study long-span data while taking into account possible regime shifts and at the same time investigate whether, if PPP is validated by the data, adjustment toward the long-run equilibrium level defined by this relationship is driven primarily by the exchange rate, by relative prices, or by both of them. Within

this framework, we are able to shed light on three hotly debated questions: whether PPP holds; whether nominal exchange rates or relative prices or both of them are responsible for responding to PPP disequilibria; and whether the half-lives of PPP deviations are consistent with standard dynamic general equilibrium open economy models with nominal rigidities.

Our results are encouraging on a number of fronts. First, our stationary MS-VECMs clearly indicate that long-run PPP holds for each of the six exchange rates examined over the last century and under each of the monetary regimes that characterize it. Second, we find that during fixed exchange rate regimes, relative prices adjust to restore deviations from long-run equilibrium, while exchange rates bear most of the burden of adjustment during flexible exchange rate regimes. Our estimated transition probabilities are consistent with the view that the relative importance of exchange rates and relative prices in restoring the long-run equilibrium level of the exchange rate varies over time and is affected by the nominal exchange rate arrangement in operation. Third, the estimated half-lives implied by our models are rather different for fixed and floating rate regimes. During fixed exchange rate regimes the effects of system-wide shocks on the PPP equilibrium relationship may last for more than two years on average with a sizable confidence interval comprising a minimum of 1.5 year and a maximum of almost 5 years. However, the corresponding half-lives during flexible exchange rate regimes are drastically shorter, presumably because the nominal exchange rate is allowed to contribute to restoring PPP. In fact, shocks will last for 0.8 years on average and the upper bound of the confidence interval is about 1 year.

Overall, we leave this study with some degree of optimism, which may be summarized in the following three points: long-run PPP may hold after all; PPP deviations are reversed more quickly under floating rate regimes, when both nominal exchange rates and relative prices are allowed to respond to shocks; the speed at which PPP deviations dissipate is consistent with the predictions of standard open economy macro theory.

Table 1. Unit root tests

		$MZ_\alpha(s_t)$	$MZ_t(s_t)$	$MZ_\alpha(\pi_t)$	$MZ_t(\pi_t)$	$MZ_\alpha(q_t)$	$MZ_t(q_t)$
Canada	(1870-2000)	-16.960	-2.566	-4.834	-1.120	-5.987	-1.609
France	(1880-1998)	-4.367	-1.433	-8.057	-1.983	-15.715	-2.921
Germany	(1880-1998)	-4.605	-1.434	-5.123	-1.526	-18.260	-3.005
Italy	(1880-1998)	-10.036	-2.239	-9.078	-2.130	-19.529	-3.041
Japan	(1885-2000)	-6.281	-1.710	-6.406	-1.733	-2.923	-0.417
UK	(1870-2000)	-6.180	-1.684	-4.408	-1.384	-18.882	-3.089

Notes: $MZ_\alpha(x)$ and $MZ_t(x)$ are modified versions of Phillips (1987) and Phillips-Perron (1988) Z_α and Z_t statistics respectively, as proposed by Ng and Perron (2001), applied to the time series $x = s_t, \pi_t, q_t$. The $MZ_\alpha(x)$ and $MZ_t(x)$ test statistics are computed using the autoregressive generalized least squares-detrended estimator to calculate the spectral density at frequency zero, s_{AR}^2 , and setting the truncation lag according to the modified BIC (MBIC) criterion as in Ng and Perron (2001). The relevant critical values for $MZ_\alpha(x)$ and $MZ_t(x)$ tests at the five percent significance level are -17.30 and -2.91 respectively (Ng and Perron, 2001, p. 1524).

Table 2. Test for common factor restrictions: linear model

	LR_{cf}
Canada	1.45×10^{-2}
France	1.34×10^{-2}
Germany	6.38×10^{-4}
Italy	9.00×10^{-8}
Japan	2.91×10^{-2}
UK	1.57×10^{-6}

Notes: LR_{cf} is the likelihood ratio test for the null hypothesis that $\beta' \Gamma(L) = G(L) \beta'$ as described in Section 2. Under the null hypothesis that the common factor restrictions are valid, the test statistic is distributed as a $\chi^2(d)$, where d is the number of restrictions. Figures reported are p -values.

Table 3. Cointegration tests

	p	L_{Max}	λ_{Trace}	χ^2	LR_{HJ}	LR_Q
Canada	2	15.99	18.85	0.007	1.75	5.82×10^{-12}
France	2	22.23	22.66	0.003	1.65	3.13×10^{-10}
Germany	2	9.16	11.60	0.442	3.24	8.66×10^{-10}
Italy	3	19.43	19.49	0.436	2.25	8.62×10^{-10}
Japan	2	12.03	13.24	0.002	2.10	6.74×10^{-10}
UK	2	13.65	13.76	0.509	3.54	1.07×10^{-9}

Notes: p is the lag order to the VAR tested for cointegration. L_{Max}, λ_{Trace} are the Johansen's LR tests based on the maximum eigenvalue and the trace of the stochastic matrix respectively. The null hypothesis tested is H_0 : no cointegration. The L_{Max}, λ_{Trace} test statistics are calculated including an unrestricted constant term in each case and, for Canada, France, Germany, Japan, also a linear trend. The critical values for L_{Max}, λ_{Trace} test statistics at the five percent significance level are 14.07, 15.41. χ^2 is the p -value for the test of the $[1, -1]$ overidentifying restrictions on the β' vector. LR_{HJ} is the fluctuation test for the null hypothesis of stability of the cointegration rank (Hansen and Johansen, 1999). The figures reported for the LR_{HJ} test are the ratios of the test statistics to the critical values at the five percent significance level; the empirical distributions of the test statistics have been calculated by Monte Carlo simulation using 100,000 iterations; $LR_{HJ} > 1$ indicates a rejection of the null of a stable cointegrating rank. LR_Q is the LR test for rank constancy proposed by Quintos (1997). The test statistics are calculated by setting the number of breakpoints $J = 1$ and the null hypothesis tested is that the number of cointegrating vectors is identical in the two subsamples split at the breakpoint. The breakpoint is identified by the recursive procedure described in Quintos (1995). The LR_Q test statistic is distributed as $\chi^2(4)$ and the figures reported are p -values.

Table 4. ‘Bottom-up’ identification procedure

	Canada	France	Germany	Italy	Japan	UK
M	3	4	3	3	4	3
$LR1$	2.65×10^{-16}	1.86×10^{-68}	3.73×10^{-46}	1.04×10^{-23}	2.90×10^{-22}	1.19×10^{-40}
$LR2$	5.35×10^{-9}	9.90×10^{-5}	3.28×10^{-8}	1.41×10^{-7}	7.76×10^{-7}	9.85×10^{-4}
$LR3$	2.04×10^{-16}	1.68×10^{-10}	1.01×10^{-14}	1.82×10^{-3}	2.14×10^{-3}	1.01×10^{-2}
$LR4$	—	—	—	2.76×10^{-4}	—	—
$LR5$	—	1.18×10^{-52}	—	—	2.35×10^{-13}	—
$LR6$	2.32×10^{-2}	1.82×10^{-20}	9.55×10^{-51}	2.03×10^{-29}	1.33×10^{-33}	1.41×10^{-2}
$LR7$	5.30×10^{-16}	5.38×10^{-76}	1.68×10^{-74}	2.83×10^{-94}	3.67×10^{-57}	5.75×10^{-53}

Notes: $LR1$ is a test statistic of the null hypothesis of no regime dependent variance-covariance matrix (i.e. $MSIA(M)$ -VECM(p) versus $MSIAH(M)$ -VECM(p)). $LR2$ is a test statistic of the null hypothesis of no regime dependent intercept (i.e. $MSAH(M)$ -VECM(p) versus $MSIAH(M)$ -VECM(p)). $LR3$ is a test statistic of the null hypothesis of no regime dependent autoregressive component (i.e. $MSIH(M)$ -VECM(p) versus $MSIAH(M)$ -VECM(p)). $LR4$ tests the null hypothesis that the model having autoregressive component of order one is equivalent to another with a higher autoregressive order (i.e. $MS(M)$ -VECM(1) versus $MS(M)$ -VECM(p)). $LR1, LR2, LR3, LR4$ are constructed as $2(\ln L^* - \ln L)$, where L^* and L represent the unconstrained and the constrained maximum likelihood respectively. These tests are distributed as $\chi^2(g)$ where g is the number of restrictions. $LR5$ is the likelihood ratio test for the null hypothesis that the MS -VECM(p) with 4 regimes is equivalent to the MS -VECM(p) with 3 regimes, and $LR6$ is the likelihood ratio test for the null hypothesis that the MS -VECM(p) with 3 regimes is equivalent to the MS -VECM(p) with 2 regimes. $LR7$ is a linearity test for the null hypothesis that the selected MS -VECM(p) is equivalent to a linear Gaussian VECM(p). p -values relative to the $LR5, LR6$ and $LR7$ tests are calculated as in Ang and Bekaert (1998). For all test statistics only p -values are reported.

Table 5. Test for common factor restrictions: regime-shifting model

	$LR_{cf}(z = 1)$	$LR_{cf}(z = 2)$	$LR_{cf}(z = 3)$	$LR_{cf}(z = 4)$
Canada	2.28×10^{-2}	9.39×10^{-3}	7.49×10^{-3}	—
France	1.51×10^{-2}	3.58×10^{-2}	9.21×10^{-3}	5.52×10^{-1}
Germany	4.86×10^{-1}	4.91×10^{-2}	9.90×10^{-3}	—
Italy	7.51×10^{-5}	3.83×10^{-2}	7.49×10^{-3}	—
Japan	1.42×10^{-1}	3.14×10^{-2}	4.56×10^{-2}	5.52×10^{-3}
UK	2.42×10^{-4}	2.26×10^{-2}	5.75×10^{-1}	—

Notes: $LR_{cf}(z = i)$ is the likelihood ratio test for the null hypothesis that $\beta' \Gamma(L \mid z = i) = G(L \mid z = i) \beta'$ conditional on regime $i = 1, 2, 3, 4$. Under the null hypothesis that the common factor restrictions are valid, the test statistic is distributed as a $\chi^2(d)$, where d is the number of restrictions. Figures are p -values are reported.

Table 6. Regime classification measure

		1870-1914	1915-1944	1945-1972	1973-2000
Canada	$\bar{p}_1$	0.008	0.203	0.052	0.200
	$\bar{p}_2$	0.848	0.434	0.678	0.078
	$\bar{p}_3$	0.064	0.364	0.270	0.722
	$RCM(3)$	2.54×10^{-1}	1.25×10^{-0}	1.05×10^{-1}	2.13×10^{-1}
France	$\bar{p}_1$	0	0.349	0.287	0.100
	$\bar{p}_2$	0	0	0.115	0.838
	$\bar{p}_3$	0.982	0.273	0.598	0
	$\bar{p}_4$	0.018	0.378	0.001	0.062
Germany	$RCM(4)$	1.71×10^{-6}	7.34×10^{-4}	2.50×10^{-8}	1.21×10^{-16}
	$\bar{p}_1$	0	0.333	0.072	0
	$\bar{p}_2$	0.999	0.388	0.810	0.004
	$\bar{p}_3$	0	0.279	0.119	0.996
Italy	$RCM(3)$	2.32×10^{-3}	9.50×10^{-2}	1.45×10^{-2}	3.98×10^{-11}
	$\bar{p}_1$	0	0.337	0.148	0.001
	$\bar{p}_2$	0.274	0.366	0.042	0.899
	$\bar{p}_3$	0.726	0.397	0.810	0.100
Japan	$RCM(3)$	4.34×10^{-3}	2.22×10^{-2}	2.64×10^{-4}	2.60×10^{-2}
	$\bar{p}_1$	0.102	0.233	0.185	0.033
	$\bar{p}_2$	0.055	0.256	0.053	0.001
	$\bar{p}_3$	0.684	0.290	0.761	0.097
UK	$\bar{p}_4$	0.159	0.220	0	0.870
	$RCM(4)$	2.63×10^{-5}	2.56×10^{-4}	7.68×10^{-9}	3.73×10^{-7}
	$\bar{p}_1$	0.065	0.301	0.203	0.767
	$\bar{p}_2$	0.934	0.359	0.673	0.053
	$\bar{p}_3$	0.002	0.340	0.125	0.180
	$RCM(3)$	4.06×10^{-1}	1.05×10^{-0}	3.09×10^{-0}	5.66×10^{-2}

Notes: $\bar{p}_j$ is the average smoothed probability relative to regime $j = 1, \dots, 4$ calculated as $\bar{p}_j = \sum_1^{T_j} p_{j,t} / T_j$, where T_j is the number of observations in the sub-period j and $p_{j,t}$ is the smoothed (ex-post) regime probability relative to regime $j = 1, \dots, M$ at time t . $RCM(M)$ is the regime classification measure as in Ang and Bekaert (2002). The statistic is calculated as $RCM(M) = 100M^M \frac{1}{T_j} \sum_{t=1}^{T_j} \left(\prod_{j=1}^K p_{j,t} \right)$ where K is the number of regimes.

Table 7. Regime-dependent adjustment coefficients

	$\alpha_1^{\Delta s_t}$	$\alpha_2^{\Delta s_t}$	$\alpha_3^{\Delta s_t}$	$\alpha_4^{\Delta s_t}$	$\alpha_1^{\Delta \pi_t}$	$\alpha_2^{\Delta \pi_t}$	$\alpha_3^{\Delta \pi_t}$	$\alpha_4^{\Delta \pi_t}$
Canada	-0.167 (0.158)	0.010 (0.010)	-0.269 (0.093)	—	0.104 (0.061)	0.042 (0.013)	-0.022 (0.048)	—
France	-0.781 (0.206)	-0.233 (0.128)	0.003 (0.003)	-0.938 (0.097)	-0.344 (0.219)	0.005 (0.012)	0.151 (0.070)	-0.129 (0.109)
Germany	-0.805 (0.478)	-0.037 (0.022)	-0.251 (0.112)	—	-0.957 (0.452)	0.029 (0.011)	0.023 (0.019)	—
Italy	-0.571 (0.566)	-0.198 (0.067)	-0.004 (0.006)	—	-0.087 (0.691)	-0.042 (0.036)	0.035 (0.010)	—
Japan	-0.385 (0.222)	0.090 (0.024)	-0.005 (0.004)	-0.312 (0.111)	-0.181 (0.213)	0.136 (0.030)	0.067 (0.030)	0.154 (0.140)
UK	-0.026 (0.008)	-0.002 (0.003)	-0.761 (0.690)	—	0.112 (0.158)	0.062 (0.022)	-0.106 (0.086)	—

Notes: $\alpha_i^{\Delta s_t}$ for $i = 1, 2, 3, 4$ denotes the error-correction coefficient relative to the spot exchange rate equation conditional to regime i . $\alpha_i^{\Delta \pi_t}$ for $i = 1, 2, 3, 4$ denotes the error-correction coefficient relative to the relative price equation conditional on regime i . Values in parentheses are asymptotic standard errors.

Table 8. Regime-conditional half lives

		Fixed regime		Flexible regime	
	z	half-life of $h(n z)$	z	half-life of $h(n z)$	
Canada	2	3.030 [1.887, 7.692]	3	0.627 [0.557, 0.772]	
France	3	2.033 [1.455, 3.497]	2	1.151 [0.981, 1.400]	
Germany	2	2.066 [1.462, 3.521]	3	0.861 [0.733, 1.048]	
Italy	3	3.226 [1.961, 9.091]	2	1.057 [0.871, 1.355]	
Japan	3	1.866 [1.381, 3.378]	4	0.649 [0.573, 0.746]	
UK	2	1.596 [1.292, 2.075]	1	0.639 [0.566, 0.735]	
average		2.303 [1.571, 4.876]		0.831 [0.714, 1.001]	

Notes: Fixed and Flexible denote fixed and flexible exchange rate regimes respectively. z denotes the regimes identified by the MSIAH(M)-VECM(p) corresponding to Fixed and Flexible (as in Table 5). The figures reported are the half-lives of the regime-dependent persistence profiles calculated as described in the Appendix with n set to 30 years. Values in brackets are the 0.025 and 0.975 quantiles of the empirical distribution of the regime-dependent persistence profiles calculated by parametric bootstrap over 100,000 replications.

A Appendix: Markov-Switching-VECM results

Table A1. MSIAH(3)-VECM(1) estimation, United Kingdom

$$\begin{aligned}
\tilde{\Gamma}_1(z_t = 1) &= \begin{bmatrix} 0.526 & 0.152 \\ (0.228) & (0.583) \\ -0.042 & 0.516 \\ (0.099) & (0.211) \end{bmatrix}; \tilde{\Gamma}_1(z_t = 2) = \begin{bmatrix} 0.002 & -0.006 \\ (0.011) & (0.014) \\ -0.026 & 0.385 \\ (0.068) & (0.089) \end{bmatrix}; \\
\tilde{\Gamma}_1(z_t = 3) &= \begin{bmatrix} 0.393 & 0.156 \\ (0.108) & (0.198) \\ -0.506 & 0.231 \\ (0.098) & (0.184) \end{bmatrix}; \\
\tilde{v}(z_t = 1) &= \begin{bmatrix} -0.012 \\ (0.007) \\ -0.009 \\ (0.003) \end{bmatrix}; \tilde{v}(z_t = 2) = \begin{bmatrix} -9.10 \times 10^{-5} \\ (1.96 \times 10^{-4}) \\ -3.94 \times 10^{-4} \\ (0.001) \end{bmatrix}; \tilde{v}(z_t = 3) = \begin{bmatrix} 0.013 \\ (0.004) \\ 0.003 \\ (0.004) \end{bmatrix}; \\
\tilde{\alpha}(z_t = 1) &= \begin{bmatrix} -0.026 \\ (0.008) \\ 0.112 \\ (0.158) \end{bmatrix}; \tilde{\alpha}(z_t = 2) = \begin{bmatrix} -0.002 \\ (0.003) \\ 0.062 \\ (0.022) \end{bmatrix}; \tilde{\alpha}(z_t = 3) = \begin{bmatrix} -0.761 \\ (0.069) \\ -0.106 \\ (0.086) \end{bmatrix}; \\
\widetilde{\Sigma}(z_1) &= \begin{bmatrix} 8.79 \times 10^{-4} & & \\ 1.57 \times 10^{-4} & 2.27 \times 10^{-4} & \\ 1.72 \times 10^{-3} & & \\ 3.55 \times 10^{-4} & 2.63 \times 10^{-4} & \end{bmatrix}; \widetilde{\Sigma}(z_2) = \begin{bmatrix} 8.44 \times 10^{-4} & & \\ 1.63 \times 10^{-4} & 2.15 \times 10^{-4} & \\ & & \end{bmatrix}; \widetilde{\Sigma}(z_3) = \\
\tilde{\mathbf{P}} &= \begin{bmatrix} 0.69 & 0.09 & 0.19 \\ 0.14 & 0.90 & 0.05 \\ 0.17 & 0.01 & 0.76 \end{bmatrix}; \quad \tilde{\xi} = \begin{bmatrix} 0.285 \\ 0.515 \\ 0.200 \end{bmatrix}; \\
\rho(A) &= 0.1043; \quad \text{LR linearity test: } 1.57 \times 10^{-72}; \quad \text{JB: } 0.382; \quad \text{RESET: } 0.254.
\end{aligned}$$

Notes: Tildes denote estimated values obtained using the EM algorithm for maximum likelihood. Figures in parentheses are asymptotic standard errors. P and ξ denote the transition matrix and the ergodic probabilities vector respectively. $\rho(A)$ is the spectral radius of the matrix A , calculated as in Karlsen (1990); it can be thought of as a measure of stationarity of the MS-VECM and stationarity requires $|\rho(A)| < 1$. The LR linearity test is a likelihood ratio test for the null hypothesis that the true model is a linear VECM(p) against the alternative of an MSIAH(M)-VECM(p). Its p -value is calculated as in Ang and Bekaert (1998). JB is the Jarque-Bera test for normality of the standardized residuals; RESET is a RESET test calculated using a third-order polynomial (Ramsey, 1969). For each of LR, JB and RESET we only report p -values.

Table A2. Relative goodness of fit

	$\overline{R}^2$ ratio		RV ratio	
	Δs_t	$\Delta \pi_t$	Δs_t	$\Delta \pi_t$
Canada	1.830	1.480	0.739	0.822
France	1.435	1.449	0.539	0.638
Germany	2.022	1.795	0.377	0.337
Italy	2.677	2.015	0.611	0.704
Japan	2.205	2.075	0.673	0.694
UK	1.773	1.288	0.750	0.881

Notes: $\overline{R}^2$ ratio and RV ratio are the ratios of the $\overline{R}^2$ and the residual variance, respectively, from each country's preferred MSIAH(M)-VECM(p) model (as selected in Table 3) to the corresponding goodness-of-fit measures for the alternative linear Gaussian VECM model.

B Appendix: Regime-dependent persistence profiles

This appendix describes the procedure we employ to analyze the response of the cointegrating equilibrium relations to particular shocks. The procedure is a generalization of the variance-based persistence profile proposed by Pesaran, Pearce and Lee (1993) and Pesaran and Shin (1996). This measure shows the effect over time of system-wide shocks on the equilibrium relations. The attractive feature of these time profiles is that they are uniquely identified since their estimation does not require prior orthogonalization of the vector of shocks. Pesaran and Shin propose the following unscaled measure of persistence:

$$H(n) \equiv V(q_{t+n}|I_{t-1}) - V(q_{t+n-1}|I_{t-1}) \quad n = 0, 1, 2, \dots \quad (\text{B1})$$

where $V(q_{t+n}|I_{t-1})$ is the variance of the long-run equilibrium vector $q_{t+n} = \beta' y_{t+n}$ conditional on the information set at time $t-1$. In the context of the VECM in equation (3), for example, the measure (B1) is given by:

$$H(n) \equiv \beta' B_n \Sigma B_n' \beta \quad n = 0, 1, 2, \dots \quad (\text{B2})$$

where $B_j = \sum_{i=0}^j \Phi_i$ ($j = 0, 1, 2, \dots$) and Φ_i are the K -dimensional matrices of the $MA(\infty)$ representation of Δy_t .²⁰ The diagonal elements of $H(n)$ are the unscaled persistence profiles or system-wide impulse responses of the cointegrating relationships $q_t = \beta' y_t$. The scaled persistence profile of the j -th cointegrating relationship $q_{jt} = \beta_j' y_t$ is given by:

$$h(n) \equiv \frac{\beta_j' B_n \Sigma B_n' \beta_j}{\beta_j' \Sigma \beta_j} \quad n = 0, 1, 2, \dots \quad (\text{B3})$$

which by construction has a unit value at the time impact when $n = 0$ and it tends to zero as $n \rightarrow \infty$ if β_j is indeed a cointegrating vector. Pesaran and Shin (1996) demonstrate that the maximum likelihood (ML) estimates of these persistence profiles are $\sqrt{T}$ consistent with a normal asymptotic distribution.

In the context of MSIAH(M)-VECM in equation (9), the scaled persistence profile (B3) can be calculated as follows:

$$h(n|z = k) \equiv \frac{\beta_j' B_n(z) \Sigma(z) B_n'(z) \beta_j}{\beta_j' \Sigma(z) \beta_j} \quad n = 0, 1, 2, \dots \text{ and } k = 0, 1, 2, \dots, M \quad (\text{B4})$$

where $B_j(z) = \sum_{i=0}^j \Phi_i(z)$ and $\Phi_i(z)$ are the K -dimensional matrices of the regime-dependent $MA(\infty)$ representation of Δy_t and $\Sigma(z)$ is the regime-dependent variance covariance matrix specified in equation (9). It is possible to gauge the precision of the estimated regime-dependent (scaled) persistence profiles by employing bootstrapping techniques. As discussed in Ehrmann,

²⁰The matrices B_j satisfy the following recursive relationship with the matrices Π_i of the VAR(p) model:

$$B_j = \Pi_1 B_{j-1} + \Pi_2 B_{j-2} + \dots + \Pi_p B_{j-p} \quad j = 1, 2, \dots, \infty$$

with $B_0 = I_m$ and $B_j = 0$ for $j < 0$. See Pesaran and Shin (1996, p. 128).

Ellison and Valla (2001, 2003), the bootstrapping technique in a Markov-switching framework is complicated by the presence of the hidden Markov-chain governing the different regimes. In fact this technique involves the generation of artificial histories of the hidden Markov-chain and then time series of the endogenous variables consistent with these histories. The full procedure requires the following four steps:

1. Create a history from the hidden regime z_t by replacing the exogenous transition matrix with its estimate $\tilde{P}$.
2. Create a history-dependent time series of endogenous variables. All parameters are replaced by their estimated values. The regime-dependent residuals are drawn from a multivariate normal distribution where the variance is calculated by means of the estimated regime-dependent variance-covariance matrices.
3. Estimate the pre-selected Markov-switching model using the series of artificial data obtained under step 2.
4. Calculate the bootstrapped estimates of the persistence profiles by using the parameter estimates obtained under step 3.

Since the artificial histories created under step 1 are typically subject to small sample bias, their estimates will not exactly coincide with those of the original data. However, by using a large number of artificial histories, we are able to obtain an accurate numerical approximation of the distribution of the estimated parameters $B_j(z)$, $\Sigma(z)$ and consequently of the scaled persistence profiles (B4).²¹

²¹For further technical details, see Ehrmann, Ellison and Valla (2001, Section 2.5).

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