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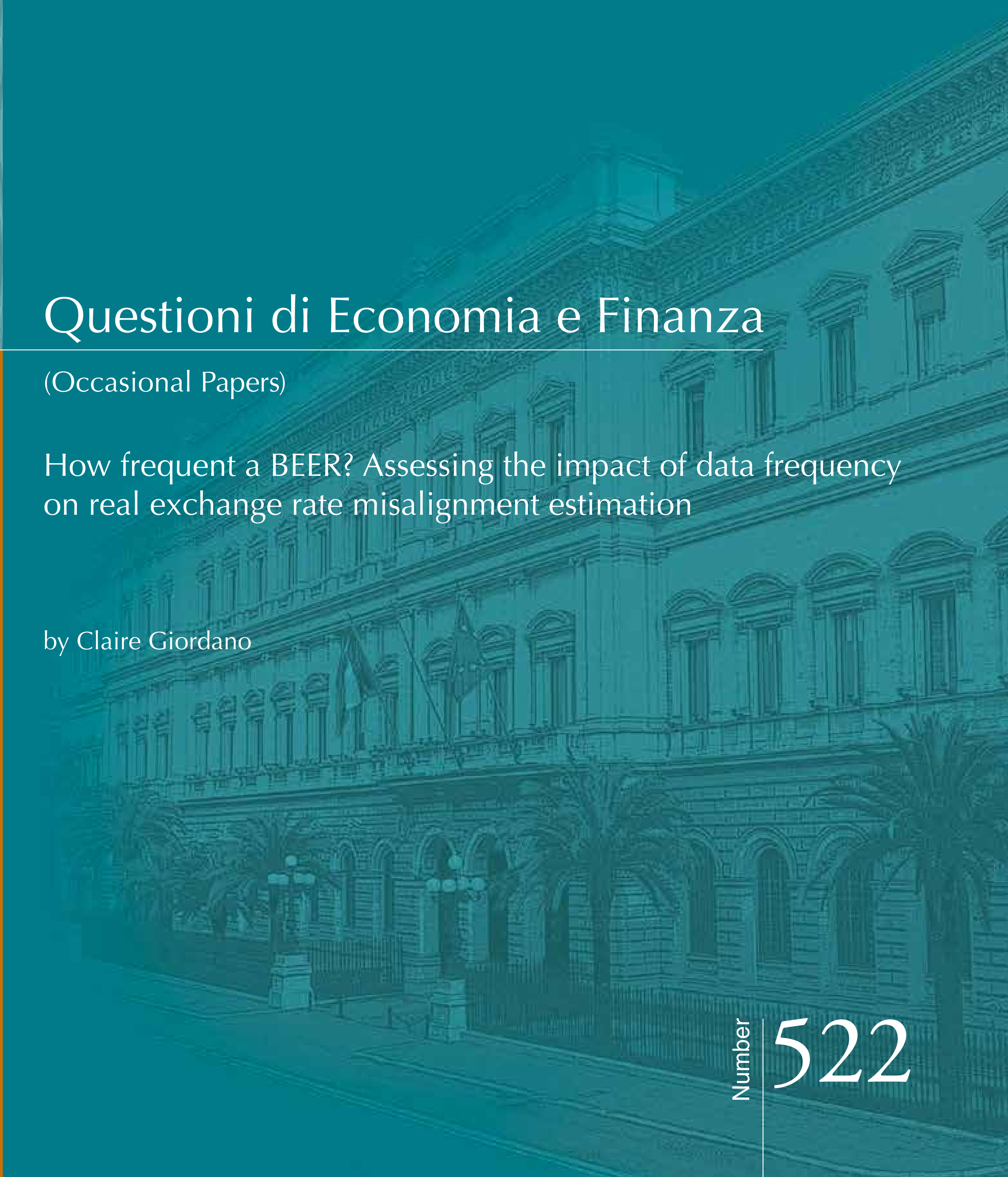

Questioni di Economia e Finanza

(Occasional Papers)

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by Claire Giordano

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HOW FREQUENT A BEER? ASSESSING THE IMPACT OF DATA FREQUENCY ON REAL EXCHANGE RATE MISALIGNMENT ESTIMATION

by Claire Giordano*

Abstract

This paper explores the robustness of Behavioural Equilibrium Exchange Rate (BEER) models – employed to estimate real effective exchange rate (REER) deviations from "equilibrium" values consistent with macroeconomic fundamentals – to the frequency (annual vs. quarterly) of the underlying data. Indeed, data frequency influences both the length of the sample period (which is typically shorter in a quarterly model) and the set of relevant fundamentals to be included in the specification, and can affect the plausibility of some of the BEER modelling assumptions, which are especially restrictive at the quarterly frequency. The paper compares REER misalignment estimates stemming from a carefully specified annual model, estimated since 1980 for 55 countries, and a comparable quarterly model, estimated since 1999, which is a variant of that currently in use at the Bank of Italy (Giordano, 2018). In the overlapping period the annualised quarterly-model misalignments are quite similar to those based on the annual model. Moreover, the in-sample power of quarterly REER misalignments in explaining subsequent, actual REER developments is found to be higher than that of the annual estimates, signalling their greater usefulness in assessing a country's external economic outlook. This paper therefore confirms the robustness of the quarterly BEER model currently employed at the Bank of Italy; moreover, it suggests that the "optimal" frequency of a BEER model depends on the use (research vs. monitoring and policy-making) one makes of the resulting measures.

JEL Classification: C54, F00, F31.
Keywords: real effective exchange rate, equilibrium exchange rate, BEER model, data frequency.
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1. Introduction[1]

The price competitiveness of a country is generally proxied by its real effective exchange rate (REER) or, using the Eurosystem terminology for the subset of euro-area countries, by its Harmonised Competitiveness Indicator (HCI). The REER is a weighted geometric average of the bilateral real exchange rates (RERs) of a country *vis-à-vis* the currency of each of its main trading partners. Seen from a different angle, the REER captures the developments in domestic prices or costs relative to the weighted average of those of its main trading partners, all expressed in a common currency.

Prices or costs are generally expressed as indices rather than as levels: they thereby only track changes in a set of prices or costs over time, and hence provide information solely on price-competitiveness dynamics, and not on how competitive a country is at a given point in time *t* relative to a benchmark, "equilibrium" level. Yet, in order to fully appraise the sustainability of a country's external position at time *t*, in addition to its price-competitiveness developments, it is necessary to assess its REER's deviation from the equilibrium value (i.e. its REER misalignment) in order to gauge whether further corrections in the REER are indeed warranted.

Various empirical models have been employed in order to operationalize the theoretical concept of this unobservable equilibrium value and to derive the resulting REER misalignment, yet no approach has been found to achieve a superior performance to the others (e.g. MacDonald, 2000; Driver and Westaway, 2004; Ajevskis et al., 2014).[2] The most common methodologies are the following three.

The first approach refers to the Behavioural Equilibrium Exchange Rate (BEER) methodology, labelled this way due to the fact that it is based on the assumption that the "behaviour" of a REER is determined by the "behaviour" of its macroeconomic drivers in the long run (Clark and MacDonald, 1998). It involves direct estimation of the reduced-form cointegrating relationship between the REER and a set of relevant economic fundamentals, which leads to the definition of the REER equilibrium value. This modelling technique does not generally have any comprehensive theoretical underpinning – since BEER models are generally defined empirically, by testing the statistical significance of many alternative potential determinants of REERs – and is a positive approach, since it is not based on any normative assumption.[3]

The second approach is the Natural Real Exchange Rate (NATREX) methodology, originally formulated by Stein (1990). It is theoretically grounded on a dynamic stock-flow model, and relies

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[2] Similar conclusions have been drawn for alternative forecasting models of REERs: see, amongst the most recent studies, Eichenbaum, Johanssen and Rebelo (2017), Cheung et al. (2018) and Ca' Zorzi and Rubaszek (2018).

[3] One possible, loose theoretical underpinning of a BEER model is Obstfeld and Rogoff's (1996) small open economy "redux" model, augmented by the transfer effect (Lane and Milesi-Ferretti, 2004) and by the presence of the government (Aguirre and Calderon, 2005), which leads to the expression of the REER as a function of productivity differentials in the tradable vs. non-tradable sectors, terms of trade, the net foreign asset position and government spending.

on a structural equilibrium concept. In particular, it defines the "natural" REER as the REER that ensures both the internal and the external equilibrium simultaneously in the long run. The internal equilibrium is achieved when the output gap is zero, that is when demand is at the level of supply potential and unemployment is at the non-accelerating inflation rate of unemployment; the external equilibrium is obtained when the current account balance is "sustainable" given a country's desired net foreign asset position, i.e. when the intertemporal budget constraint is satisfied. Although there have been some attempts to measure the structural model underlying the NATREX (e.g. Gandolfo and Felettigh, 1998; Siregar and Rajan, 2006), this approach often boils down to estimating a reduced-form equation. In this case the main difference between the BEER and the NATREX is only that the latter is more explicitly theory-based (Stein, 2006). Again, similarly to the BEER methodology, the NATREX model adopts a positive approach.

The third approach is the Fundamental Equilibrium Exchange Rate (FEER) methodology, advocated by Wren-Lewis (1992) and Williamson (1994). Similarly to the NATREX approach, the FEER is the REER that simultaneously attains internal and external balance. In its most popular applications (Isard, 2007; Lee et al., 2008; Cline and Williamson, 2010), the FEER method is based on a partial equilibrium model and, in particular, on the computation of the REER adjustment required to close the gap between the cyclically-adjusted current account and the "current account norm", which represents an optimal value of the current account over a medium-term horizon. The norm is either set in a normative manner or is derived from "behavioural" reduced-form regressions that estimate an equilibrium relationship between the current account and a set of plausible fundamentals. The calibration of the required REER adjustment is, however, highly sensitive to the assumptions made concerning both exchange-rate pass-through coefficients and price elasticities of trade (Schnatz, 2011). The FEER model can be reconciled with the NATREX approach by estimating a target level for the net foreign asset position rather than for the current account balance.

These approaches, far from being opposed to each other, are complementary, both because they are all based on restrictive (albeit different) assumptions and because they may be interpreted as assessing equilibrium exchange rates over different time horizons (Bénassy-Queré, Béreau and Mignon, 2010). Indeed, the FEER, for instance, may be considered as corresponding to a medium-run concept of equilibrium, whereas the BEER to a long-run concept.

Currently, the IMF is, to our knowledge, the only institution that provides estimates of REER misalignments based on all three approaches, in the context of its External Balance Assessment (Phillips et al., 2013; Cubeddu et al., 2019), via its current-account model, its REER regressions and its external sustainability approach (where the equilibrium REER is defined as the rate that closes the gap between a country's actual current account balance and the balance that would stabilise the net foreign asset position of the country at a benchmark level). These three methods indeed broadly correspond to FEER, BEER and NATREX models, respectively.

This paper focuses solely on one methodology, namely the BEER approach. Yet, even when focusing on one method only, many modelling choices arise, concerning the selection of the country sample, the estimation period, the data frequency, the weights and the deflator employed to construct the REERs, the numeraire currency against which to express the bilateral RERs underlying the REERs, the relevant set of economic fundamentals, the appropriate proxies of these economic fundamentals, as well as several econometric aspects (in particular, linked to accounting for country heterogeneity and tackling cross-sectional dependence, as shall be clarified below).

Few studies amongst the many that estimate BEER models (a selection of which are reported in Table A1 in Annex A) discuss and systematically address these manifold methodological issues. Recent exceptions are the following. Bussière et al. (2010) focuses largely on the appropriate econometric techniques to be employed in estimating BEER models. Bénassy-Queré, Béreau, and Mignon (2010) mainly addresses the issue of the correct measurement of the relevant economic fundamentals, namely the Balassa-Samuelson effect, and of the choice of the numeraire country. Couharde et al. (2017) accounts for country sample selection, the choice of different economic fundamentals and the REER weighting procedure. Adler and Grisse (2017) is centred on the selection of economic fundamentals, the country sample and the estimation methodology. Fidora, Giordano and Schmitz (2018) and Giordano (2018) adopt various price and cost indicators to construct the REERs, as well as employing alternative Balassa-Samuelson proxies, carefully selecting the set of relevant explanatory variables and using different estimation techniques. Finally, Fischer (2019) tests alternative estimation methods.

Adding to these studies, this paper is concerned with the robustness of equilibrium REERs and of the corresponding misalignments according to a BEER approach along one dimension only, namely the choice of the data frequency (i.e. annual vs. quarterly). Various methodological aspects depend on this choice, such as the selection of the sample period and of the relevant set of economic fundamentals. All other modelling choices are held fixed between the annual and the quarterly models, namely the country sample, the range of price and cost indicators employed to compute the REER, the numeraire currency, and the econometric technique used for estimation.

A priori, quarterly models have two main advantages. First, the higher number of observations relative to an annual dataset implies that quarterly models are richer information-wise and may be estimated more efficiently (as in any economic context; e.g. Silvestrini and Veredas, 2008). Second, estimating a BEER model at quarterly frequency is of paramount importance for monitoring, policy-making and early-warning purposes, since it allows tracking infra-annual imbalances in a timely manner (Giordano, 2018).

Conversely, the main drawback of adopting higher frequency data is thinner data availability, and therefore the restriction of the estimation window to a shorter time-span, covering only the most recent years. This limitation, in turn, exacerbates two significant shortcomings underlying the BEER methodology.

The first shortcoming concerns the need to include country fixed effects amongst the regressors included in the BEER model, due to the fact that the dependent variable, the REER – generally based on price or cost indices – is an index number (Fischer and Hossfeld, 2014; Adler and Grisse, 2017; Cubeddu et al., 2019). As mentioned earlier, REERs thus computed provide information on the evolution of the relative purchasing power of two currencies relative to the base year of the employed price or cost index, and no information on the actual relative purchasing power level. The inclusion of country fixed effects is therefore a means to account for country-specific price levels in the base year. Yet, by including fixed effects, the predicted (i.e. equilibrium) REERs are by construction on average equal to the long-run actual REER means, or in other terms each country's regression residuals – and thus REER misalignments – are forced to average to zero over the sample period. Hence, "persistent" deviations of REERs from the equilibrium value in the estimation sample are by construction not envisaged in BEER models.[4]

The second drawback of the BEER methodology is that it implicitly assumes that the economic fundamentals – against which the actual REER is appraised – are at their equilibrium values. This is less likely to be the case for short time-spans.

Moving to annual data implies being able to significantly extend the sample period (in this paper, since 1980) and therefore making both the assumptions of zero misalignments and of economic fundamentals at their equilibrium values on average over the estimation horizon more palatable. Yet, one may argue that even (nearly) forty years may not be enough for these hypotheses to be grounded (e.g. Taylor, 2002, which employs over a century of data). Therefore, time coverage will always be an issue, also considering the fact that the analysis of lengthy periods may negatively affect data quality, as well as raise the issue of how to handle structural breaks over time. In our view, forty years is a good compromise between data reliability and sufficient length of the sample period; moreover it is a relatively long time span when compared to the existing BEER model literature (again, see Table A1).

This paper first describes and estimates an annual BEER model since 1980 for a vast sample of advanced and emerging countries, using recently developed techniques both to select the relevant economic fundamentals (Bayesian model averaging) and to estimate the model (common correlated effects mean group estimation). Particular care is also adopted in constructing appropriate proxies of the relevant economic fundamentals, in particular the Balassa-Samuelson effect and trade costs. A comparable quarterly BEER model – which differs from the annual model only due to the underlying data frequency – is also put forward; based on Giordano (2018), which is the model currently in use at the Bank of Italy, it spans from 1999 due to quarterly data availability. Next, the annual HCI misalignment estimates obtained for the four main euro-area economies as a relevant case-study are compared with the quarterly-model results in order to gauge any significant difference. Finally, so as to explore the usefulness of the two sets of estimates in appraising a country's external economic outlook, the paper assesses the in-sample performance of annual vs. quarterly REER/HCI misalignments in explaining subsequent REER/HCI developments in all countries, by adopting the regression framework in Abiad, Kannan and Lee (2009).

The main findings are the following. The economic fundamentals to be included in the BEER model, on the basis of their statistical significance, do indeed differ according to the data frequency: whereas demographics and capital accumulation are statistically significant at annual frequency, they do not matter at a quarterly frequency; conversely, real interest rates are significant quarterly determinants, whereas they are irrelevant at an annual frequency. HCI misalignment estimates for the four main euro-area countries stemming from the annual and quarterly models are found to be generally comparable; some significant discrepancies emerge only for Spain and Germany in some years, plausibly due to substantial changes in their (relative) investment rate and old-age dependency

[4] The exclusion of the possibility of modelling persistent misalignments in BEER models may be particularly unappealing in certain periods. For example, as cited by Romer (2012, p. 646), "[a]nother message of the [recent global financial] crisis concerning financial markets is that there are limits to the forces bringing asset prices in line with fundamentals", a claim which can also be applied to REERs.

ratio, which only the annual model captures. Finally, quarterly misalignments are found to have a higher explanatory power for all countries' subsequent REER/HCI movements. Hence, in spite of the more binding assumptions underlying a quarterly as opposed to an annual BEER model, this paper validates the use of quarterly models, and in particular of that currently in use at the Bank of Italy. Moreover, one can conclude that no "optimal" frequency of BEER models exists. The chosen frequency should depend on the pursued aim: for example, for research purposes (e.g. studies on the link between REER misalignments and economic growth) annual estimates, available for nearly forty years, are more appropriate; the more timely quarterly estimates are instead better geared for monitoring, early-warning and policy-making purposes.

The remainder of this paper is structured as follows. Section 2 describes the general set-up of a BEER model. Section 3 describes the specification and the estimation of the annual BEER model. Section 4 summarises the comparable quarterly BEER model. Section 5 first compares estimates of HCI misalignments for the four main euro-area economies according to the two models; it next evaluates, for all the countries in the panel, the explanatory power of annual vs. quarterly REER/HCI misalignments in explaining subsequent REER/HCI developments. Section 6 concludes.

2. The general set-up of a BEER model

In a nutshell, a BEER model estimates the long-run reduced-form relationship between bilateral real exchange rates (RERs) or real effective exchange rates (REERs), on the one hand, and key macroeconomic fundamentals, on the other hand (Clark and MacDonald, 1998). The in-sample predictions of the model provide estimates of equilibrium values; the percentage-point difference between the actual and the equilibrium R(E)ERs is labelled as the R(E)ER misalignment.

The starting point of a BEER model is the basic arbitrage condition of uncovered real interest parity, which holds under perfect capital mobility, free trade and rational expectations, according to which:

(1) $E_t(rer_{t+1}) - rer_t = -(r_t - r_t^*)$

where rer_t and rer_{t+1} are the RER of a given country at time t and $t+1$, respectively, defined such that an increase refers to RER appreciation, r_t and r_t^* are the domestic and foreign real interest rates and E_t denotes the expected value at time t. This relationship postulates that the expected return on a domestic riskless short-term interest rate-bearing security and a corresponding foreign security have the same expected return, when returns are expressed in the same currency.

By rearranging the terms in equation (1), the observed RER in time t is expressed as a positive function of the expected value of the next-period RER (which coincides with the "equilibrium" RER in steady state, i.e. in the absence of any further shock to the domestic and foreign economies) and of the current real interest-rate differential:

(2) $rer_t = E_t(rer_{t+1}) + (r_t - r_t^*)$

Clark and MacDonald (1998) assume that the unobservable expected future value of the RER is determined by a vector of economic fundamentals, so the actual RER ultimately depends on these drivers and on the real interest-rate differential. The operationalization of equation (2) therefore requires the definition of the RER, on the one hand, and the selection of economic fundamentals, on the other.

In practice, both RERs and REERs have been used as the dependent variable in BEER models. Since the dependent variable is a bilateral or multilateral concept, also the economic fundamentals on the right-hand side need to be expressed either bilaterally (i.e. relative to a numeraire country) or multilaterally (i.e. against a weighted average of the set of trading partners of a given country) in order to guarantee multilateral consistency of the resulting REER misalignments. The implications of the choice of the dependent variable and of the numeraire country will be better discussed in Section 3. For the sake of brevity, in the remainder of this section we will only refer to RERs, but all claims also apply to REERs.

Another choice concerning the dependent variable is how to deflate the RER. The generally employed price or cost indices – such as the consumer price index (CPI), the GDP deflator, the producer price index (PPI) or the unit labour cost of the total economy (ULCT) – require the inclusion of country fixed effects in the BEER model, as discussed in Section 1. They therefore lead to the assumption of zero average misalignment over the sample period, which may be implausible over short time spans, typical of quarterly models. This issue has been tackled, for instance, by employing a price level, the purchasing power parity (PPP) rate, as a deflator (e.g. Adler and Grisse, 2017; Giordano, 2018; Cubeddu et al., 2019), which can lead to the exclusion of country fixed effects. However, this choice too is not flawless. On the one hand, if one drops the fixed effects, the explanatory variables expressed as index numbers (such as terms of trade) also need to be excluded from the model in order to obtain reliable estimates, although they may be relevant determinants of REERs. On the other hand, PPPs suffer from measurement issues that do not affect price indices (Deaton and Heston, 2010). This confirms the usefulness of comparing REER misalignment estimates based on alternative (index and level) deflators, as done in this paper.

Turning to the economic fundamentals, one of the most popular explanations of the deviations of RERs from equilibrium is due to Balassa (1964) and Samuelson (1964), which posited that relative prices of non-traded and traded goods are inversely related to the relative productivity in the two sectors, under a set of assumptions, including free labour mobility across sectors. In a partial equilibrium setting, a rise in productivity in the tradable sector entails an increase in wages in the same sector, yet also bids up wages in the non-tradable sector. Since the wage increase in the non-tradable sector is not accompanied by productivity gains, this leads to a higher general price level, which in turn implies a RER appreciation.

Whereas the Balassa-Samuelson (hereon BS) model assumes that the RER depends entirely on supply factors, demand-side variables are also typically considered in BEER models, of which the most frequently employed are the following.[5]

[5] Indeed, as noted by Froot and Rogoff (1995), demand factors can have an effect on the relative price of non-traded goods if one of the following assumptions of the BS model is relaxed: perfect competition in goods markets, perfect mobility of labour between the two production sectors, international mobility of capital, law of one price for traded goods and constant returns to scale in the two sectors.

First, international trade policy may have an effect on RERs. Lower trade barriers can lead to RER depreciation via a fall in domestically produced goods' prices, in turn due to heightened competition or to cheaper intermediate inputs (Goldfajn and Valdes, 1999; Ricci et al., 2013).

Second, terms of trade are often included in BEER models as they are generally considered a source of fluctuations of RERs that can be considered as exogenous, since few countries exert significant market power on export and import prices (Lane and Milesi Ferretti, 2004). In particular, an improvement in terms of trade, e.g. an increase in export prices relative to import prices, should lead to a positive income or wealth effect in the domestic economy. The ensuing rise in domestic demand for and the reduction in the foreign supply of non-tradables increases domestic non-tradable prices and therefore leads to a RER appreciation (Neary, 1988). If there is a home bias in the tradables consumption basket, this will reinforce the real appreciation following a rise in the terms of trade (Lane and Milesi-Ferretti, 2002).

An exclusive focus on the terms of trade however neglects the impact of international transfers of resources between countries, for instance connected to interest payments, on the relative price of non-tradables. Therefore, in addition to terms of trade, a third factor, namely net foreign assets (NFAs), are often included in BEER models, and operate through the (endogenous) relative price of non-tradables (Obstfeld and Rogoff, 1996; Lane and Milesi-Ferretti, 2004). In particular, in intertemporal optimizing models the transfer effect can operate in the presence of a home preference for domestic tradables or though the impact of wealth effects on labour supply. In the former case, a transfer from the foreign to the home country implies an increase in demand for home goods and a rise in their relative price. In the latter case, a similar transfer will increase domestic wealth, lower labour supply and the supply of tradables, raising their relative price. In the long run countries will be bound by their intertemporal budget constraint: lenders will demand repayment of their loans and borrowers will have to reimburse their debts. Hence, countries with significant external liabilities eventually need to run trade surpluses in order to service the interest payments due, and thus their RER needs to depreciate; conversely, a positive NFA position enables a country to run persistent trade deficits, which in turn, all else equal, requires an appreciated RER. The (conditional) correlation between RERs and NFA is therefore expected to be positive.

Fourth, final government consumption can positively affect the RER through a composition effect (Froot and Rogoff, 1992; Obstfeld and Rogoff, 1996; Hinkle and Montiel, 1999): because government consumption tends to fall disproportionately on domestic non-tradables, the RER tends to rise as a result of a surge in this demand component. However, excessive government consumption, and therefore spending, may cast doubt on the sustainability of fiscal policy and undermine the confidence in a country's currency, leading to RER depreciation (Frenkel and Mussa, 1985; Melecký and Komárek, 2007). The expected correlation between government consumption and the RER is therefore *a priori* ambiguous.[6]

[6] Galstyan and Lane (2009) underscore how an increase in public investment, as opposed to consumption, may also affect the RER, yet with an a priori ambiguous impact. By delivering productivity gains in the tradable sector a rise in public investment may indeed generate real appreciation through the BS mechanism, but if it disproportionately raises productivity in the non-tradable sector, it may actually lead to real depreciation. Herein, however, we only focus on government consumption, as is standard in the BEER model literature, due to data availability.

Fifth, demographics may affect the RER (e.g. Giagheddu and Papetti, 2018). Under the life-cycle hypothesis, lower labour participation or a higher old-age dependency (OAD) ratio can lead to a more appreciated RER. In particular, the elderly consume more non-traded services relative to working-age people, implying an increase in overall demand for those goods in the presence of population ageing (Groneck and Kaufman, 2017). At the same time, the old-age population has lower saving rates than younger cohorts, such that aggregate savings of an ageing society decline (Higgins, 1998; Yoon, Kim and Lee, 2014) and aggregate consumption increases, again biased towards non-tradable goods. If the additional demand for non-traded services of an ageing society is not fully met by higher supply, the relative price of non-tradables increases and the RER appreciates. Another channel that links demographics and the RER is the wealth effect explored in the overlapping-generations model in Aloy and Gente (2009). With selfish agents, aggregate wealth depends on the proportion of new-born individuals in the total population. Since new-borns hold zero financial wealth, a fall in the birth rate leads to an increase (decrease) in wealth per capita in the case of a creditor (debtor) country. Consumption therefore increases (decreases) too, causing a RER appreciation (depreciation).

Sixth, the investment rate can proxy for technical progress (Bussière et al., 2010), which can lead to productivity rises and therefore to a RER appreciation. However, given investment's high import content, it may also affect the trade balance negatively, with an opposite impact on the RER.

Seventh, the monetary policy stance can be captured by the real interest rate (e.g. Adler and Grisse, 2017). An (unanticipated) increase in real interest-rate differentials should give rise to capital inflows and therefore to a RER appreciation.

Finally, (de-trended) credit to the private sector as a share of GDP has been employed as an indirect indicator of financial excesses (Cubeddu et al., 2019). The latter may cause demand booms, leading to RER appreciation.

Once the specification of the BEER model has been pinned down, the model may be estimated in order to derive its fitted values, i.e. the RER equilibrium values. Using a panel dataset on a wide sample of economies, as opposed to single country regressions, is now standard practice when estimating BEER models. Exploiting both the time and the cross-section dimensions indeed increases sample size and raises the power of statistical tests, as well as achieving more robust estimates.

However, using a unique panel equation for calculating equilibrium exchange rates relies on the very strong assumption that the same behaviour of economic fundamentals, and therefore, RERs applies to all countries, which often include both advanced and emerging economies. To some extent, this is a desirable property: the economic fundamentals that drive RERs in the long term should be the same across countries, especially since, looking forward, emerging economies should behave more like advanced economies. In other terms, estimating a single equilibrium exchange rate equation for all countries allows smoothing the impact of individual countries' transitional dynamics (Bénassy-Quéré, Lahrèche-Révil and Mignon, 2008; 2009). However, the exact relation between the dependent variable and each of its drivers may differ across countries. Allowing for various forms of country heterogeneity in the panel estimation procedure is now possible owing to the latest-generation panel cointegration techniques (e.g. Hlouskova and Osbat, 2009; Hossfeld, 2010), as we later explain and apply.

Another technical issue that a panel setting raises is that of cross-section dependence (CSD; see, for example, Bussière et al., 2010; Chudik, Pesaran and Tosetti, 2011). Indeed, when both the number of countries and time periods are large, cross correlation of errors may emerge, due to omitted, unobserved common effects to all countries, such as global demand and supply shocks. In the specific BEER model setting, the case for CSD is especially compelling, given the fact that all explanatory variables are expressed relative to a numeraire country. Conventional panel estimators such as fixed or random effects can result in misleading inference and even inconsistent estimators, depending on the extent of CSD and on whether the source generating the CSD (such as an unobserved common shock) is correlated with the regressors (Chudik and Pesaran, 2015). Recently developed panel cointegration techniques – which we employ in our models, as later shown – can also correct for this potential bias.

After having outlined a general BEER model framework, we now move on to describe first the annual and then the quarterly models employed in this paper.

3. The annual BEER model

The annual BEER model is estimated since 1980 for 55 euro and non-euro area countries, listed in Table 1. The economies considered include both advanced and emerging countries, accounting for over 90 per cent of global GDP in 2017, a relatively large country coverage in comparison with that of the other BEER model studies reported in Table A1. For few economies, such as countries of the former Soviet Union, Yugoslavia and Czechoslovakia, series of equilibrium exchange rates and misalignments start at later dates.

Table 1. The list of countries

Euro area	Other advanced economies	Emerging economies
Austria (AT)*	Australia (AU)*	Algeria (DZ)*
Belgium (BE)*	Canada (CA)*	Argentina (AR)
Cyprus (CY)*	Czech Republic (CZ)*	Brazil (BR)
Estonia (EE)*	Denmark (DK)*	Bulgaria (BG)*
Finland (FI)*	Hong Kong (HK)*	Chile (CL)
France (FR)*	Iceland (IS)**	China (CN)*
Germany (DE)*	Israel (IL)	Croatia (HR)*
Greece (GR)*	Japan (JP)*	Hungary (HU)*
Ireland (IE)*	Korea, Republic of (KR)*	India (IN)
Italy (IT)*	New Zealand (NZ)	Indonesia (ID)
Latvia (LV)*	Norway (NO)*	Malaysia (MY)
Lithuania (LT)*	Singapore (SG)*	Mexico (MX)
Luxembourg (LU)**	Sweden (SE)*	Morocco (MA)
Malta (MT)*	Switzerland (CH)*	Philippines (PH)
Netherlands (NL)*	United Kingdom (GB)*	Poland (PL)*
Portugal (PT)*	United States (US)*	Romania (RO)*
Slovakia (SK)*		Russian Federation (RU)
Slovenia (SI)*		South Africa (ZA)
Spain (ES)*		Thailand (TH)
		Turkey (TR)*

* Narrow country sample for which ULCTs are available.
** Countries for which PPIs are not available.

3.1 The dependent variable

As in Fischer and Hossfeld (2014) and Mancini Griffolo et al. (2015), the dependent variable of our BEER model is the bilateral RER, as opposed to the multilateral REER. This choice, which is discussed more thoroughly in Annex A, presents the advantage that the RER captures relative prices in a cleaner fashion, since it is unaffected by changes in trade weights, which in turn can be endogenous to exchange-rate variations as they modify the relative value of trade flows across partners. More importantly, the approach *per se* ensures the multilateral consistency of estimated misalignments, given that the effective equilibrium value of each currency can be calculated as a (trade-)weighted average of its bilateral exchange rate equilibria, as explained later in detail.

In particular, the selected dependent variable is the bilateral RER relative to the US dollar. While a number of authors find that the choice of the numeraire currency does not significantly affect the computation of REER equilibrium levels and misalignments (e.g., Bénassy-Queré et al., 2004; Bénassy-Queré, Béreau and Mignon, 2008 and 2009; Fidora, Giordano and Schmitz, 2018), Hlouskova and Osbat (2009) argues that – although in a bilateral estimation set-up the choice of the numeraire will not qualitatively impact the coefficient estimates – the aggregation of bilateral misalignments into the effective misalignment will lead to estimates that are affected by the effective misalignment of the numeraire currency at all points in time. The study hence suggests using time fixed effects in order to control for the misalignment of the numeraire. In this paper we at least partly account for this potential bias by adding cross-section averages of both the dependent and explanatory variables to correct for CSD, as will be explained in Section 3.2. We also conduct some robustness checks on this issue, discussed in Section 3.4.

In order to obtain the dependent variable, yearly average nominal exchange rates are deflated by one of the following indices: i) the CPI, ii) the GDP deflator, iiii) the PPI, iv) the ULCT; or by the following price level: v) the PPP rate[7] (Table 2). The PPI is available only for 53 countries (in particular, Luxembourg and Iceland are not available in the employed Bank of Italy dataset; see Felettigh and Giordano, 2018) and the ULCT is only available for a narrow sample of 38 countries (again see Table 1), and only as of 1995.

Indeed, in spite of the ongoing debate on the topic (e.g. Chinn, 2006; Giordano and Zollino, 2016; Ahn, Mano and Zhou, 2017), there is no consensus on the optimal price or cost to employ in the construction of RE(E)Rs, which makes it necessary to consider a range of alternative indicators. As seen in Table A1, however, BEER models have mainly included relative CPIs or PPPs. To our knowledge, Fidora, Giordano and Schmitz (2018) and Giordano (2018), as well as this paper, are the only attempts to consider such a wide range of price and cost indicators. Moreover, the importance of double-checking results based on price indices as opposed to levels has already been stressed in Section 2.

[7] In this paper, as seen in Table 2, we use the PPP USD rates sourced from IMF-WEO which cover the whole time span under study herein, but rescale the series by using the time-varying PPP rate for the US sourced from the Penn World Tables, for the years available, in order to preserve both the cross-section and the time-series dimensions (see Giordano, 2018 for a discussion of this issue, as well as Cubeddu et al., 2019 for an alternative way of constructing relative PPPs).

Table 2. The variables in the annual BEER model: sources and details

	Notes	*Sources*
1. Dependent variable		
Nominal exchange rates	Relative to the US dollar. Deflated with one of the following five indicators.	IMF WEO
CPI		ECB, IMF WEO, BIS
PPP	The IMF WEO series are rescaled with the time-varying PPP series of the United States sourced from the Penn World Tables, for the years available (until 2014).	IMF WEO; Penn World Tables
GDP deflator		ECB, IMF WEO and IFS, World Bank WDI
PPI	Available for 53 countries.	Bank of Italy
ULCT	Available for 38 countries and since 1995.	ECB
2. Explanatory variables		
GDP per capita	PPP-based.	IMF WEO
GDP per worker	Computed as the ratio of PPP GDP to headcount employment.	IMF WEO and IFS, World Bank WDI
Phi-ness indicator	The bilateral indicator, sourced from ESCAP-WB Trade Cost Database, is weighted with the trade weights employed to construct the REER. Available for 1995-2015. For the remaining years it is interpolated with the trade openness variable.	ESCAP-WB Trade Cost Database, ECB, IMF-WEO
Trade openness	Computed as the sum of current-price total exports and total imports as a share of current-price GDP.	IMF-WEO
Terms of trade	Terms of trade index when available, otherwise computed as the ratio of the export unit value index to the import unit value index.	IMF WEO and IFS, World Bank WDI
Government consumption	Computed as the ratio of current-price government consumption to current-price GDP.	IMF WEO and IFS, World Bank WDI, OECD
Investment rate	Computed as the ratio of current-price gross fixed capital formation to current-price GDP.	IMF WEO and IFS, World Bank WDI
Old-age dependency ratio	In percentage of working-age population.	World Bank WDI
Net foreign assets	In US dollars.	Lane and Milesi-Ferretti (2018) and IMF IFS
Real interest rate	Computed as the nominal interest rate deflated by the CPI inflation rate.	IMF WEO and IFS, World Bank WDI, BIS
Private sector credit	Computed as the ratio of bank credit to the private sector in US dollars to current-price GDP in US dollars, detrended via the Hodrick-Prescott filter.	BIS, IMF WEO
3. Robustness		
CPI-to-PPI ratio	Computed as the ratio of CPI and PPI. Available for 53 countries.	ECB, IMF WEO, BIS, Banca d'Italia
GNP per capita	PPP-based. Available since 1990.	World Bank WDI
Young-age dependency ratio	In percentage of working-age population.	World Bank WDI
Aging speed	Computed as the change in the total old-age dependency rate twenty years ahead relative to the current period.	World Bank WDI
Labour participation rate	In percentage of population.	IMF WEO, World Bank WDI
Railroad density	Computed as the ratio of railways (route-km) to land area (squared km). Available until 2016.	World Bank WDI
4. Other		
Trade weights	3-year average import and double export weights vis-à-vis 54 trading partners.	ECB

3.2 The economic fundamentals

As discussed in the previous sections, in the BEER model literature there is no unambiguous prior theory for the selection of economic fundamentals; those proposed as RER drivers in previous studies cannot be excluded *a priori*, if not due to data availability issues. Yet, it is not necessarily optimal to include all possible fundamentals in a BEER model. There is indeed a trade-off between using potentially redundant variables, which result in less precise coefficient estimates, and a potential omitted variable bias, which could distort estimates if the omitted variable is correlated with the other regressors. Our first step is therefore to construct the explanatory variables suggested by the literature and for which data are available for our panel of country-years. In the second step we handle model uncertainty econometrically.

The details on the construction of the various variables and on their data sources are reported in Table 2. Since the dependent variable is expressed relative to the US dollar, all explanatory variables are always expressed relative to their US counterpart.

To empirically investigate the BS effect, in principle data on sectoral productivity should be employed in order to construct tradable vs. non-tradable productivity differentials across countries. Due to data availability, this is not feasible for our panel.[8] Yet, as shown in Lothian and Taylor (2008), if productivity in the non-tradable sector is constant across countries, total-economy productivity differentials may be employed as a proxy. This assumption is strong, yet there is evidence for many countries over long time spans that productivity growth in private service sectors, a proxy of the non-tradable sector, is significantly slower than that of sectors open to trade, and often close to zero across countries (e.g. Timmer, Inklar and O'Mahoney, 2010; Giordano and Zollino, 2019). Against this background, the BS effect is often simply proxied by relative GDP per capita (e.g. Aguirre and Caldéron, 2005; Hassan, 2016; Couharde et al., 2017), a practice we too follow. However, in addition to this measure, we also adopt relative labour productivity.

In order to measure international trade restrictions the BEER model literature has mainly employed, as an inverse proxy, the openness to trade ratio, i.e. the sum of exports and imports as a share of GDP, under the assumption that countries with more liberal trade regimes have higher trade volumes, *ceteris paribus*. The issue with the trade-openness indicator is that it may be simply capturing the size of the economy, since large countries are generally more closed than small economies. And since both the size of the non-tradable sector and the degree of exogeneity of terms of trade are related to country size (Lane and Milesi-Ferretti, 2004), this proxy may be capturing different mechanisms to the trade policy channel. Moreover, contemporaneous exchange-rate fluctuations may affect the trade-openness indicator, leading to a reverse causality issue especially in an annual BEER model. For the years for which it is available, we therefore employ a more direct measure of trade costs, namely the so-called "phi-ness" indicator of trade, put forward by Head and Ries (2001), Baldwin et al. (2003) and Head and Mayer (2004). The core idea of this micro-founded measure is that the trade cost parameters of a theoretical gravity equation may be derived from observable bilateral trade data. This indicator covers all costs involved in trading goods internationally with another partner relative to those involved in trading goods domestically, as

[8] As discussed more thoroughly in Giordano (2019), in order to compare sectoral levels of productivity across countries, sector-specific PPPs are necessary. These are not available if not for some benchmark years.

explained more rigorously in Annex A. It therefore captures a wide range of trade cost components, such as transportation costs and tariffs, linguistic and cultural barriers, informational costs and bureaucratic red tape. To our knowledge, we are the first to employ this indicator in a BEER model context. This measure – sourced from the ESCAP-WB Trade Cost Database, then trade-weighted so as to move from a bilateral to an effective indicator *vis-à-vis* all trading partners and then expressed relative to that of the US – is however available only for the years 1995-2015. Yet, as shown in Figure A1 of Annex A for the four main euro-area countries, the correlation between the phi-ness indicator and openness to trade is relatively strong (approximately -0.7). The former variable is therefore spliced with (the inverse of) the latter measure in order to fill the missing years.

The other economic fundamentals are constructed in a standard fashion based on national account or financial data retrieved from various sources (IMF WEO and IFS, OECD, BIS, World Bank WDI). In particular, demography is captured by four alternative proxies: the labour participation rate; the OAD ratio; the young-age dependency ratio; the projected aging speed, computed as the change in the total OAD ratio twenty years ahead relative to the current period, as in Cubeddu et al. (2019). The construction of the NFA data is thoroughly documented in Lane and Milesi-Ferretti (2018), to which we refer; we update their dataset with recent figures from the IMF IFS. The private-credit to GDP ratio is de-trended using traditional filtering techniques in order to gauge any "financial excesses", as is standard in the financial economics literature.[9]

Having constructed the dataset, in order to select the economic fundamentals to be actually included in the baseline BEER model, we next estimate regressions for all possible combinations of the explanatory variables, using Bayesian model averaging (BMA) techniques. BMA involves the estimation of all possible sets of variables, assigning to each set a posterior model probability of being "true" based on Bayesian inference, and then the computation of a weighted average of all the estimates for a given coefficient in a statistically optimal way. Inference can then be based on the whole universe of candidate models, by considering not only the uncertainty associated to the coefficient estimate conditional on a given model, but also the uncertainty of the coefficient estimate across different models. A formal derivation of BMA is provided in Annex A. To our knowledge the other studies applying this formal "horse-race" method to appropriately specify a BEER model are only a handful, namely Bussière et al. (2010), Du, Wei and Xie (2013) and Adler and Grisse (2017).[10] This methodology aims at detecting systematic empirical regularities, whereas precisely identifying causal relationships would require a more structural approach that cannot handle this amount of data, and which anyhow goes well beyond the scope of BEER modelling.

In practice, via BMA different subsets of so-called "auxiliary" regressors (i.e. the potential determinants of the RER in our case) can be excluded from the model to improve, in the mean squared error sense, the efficiency of the "focus" parameter estimates (i.e. the coefficients attached to the

[9] Specifically, as in Cubeddu et al. (2019), a Hodrick-Prescott (Hodrick and Prescott, 1997) filter is applied to the credit-to-GDP ratio with a large penalty parameter that takes into account the fact that financial cycles have longer duration than real business cycles. By applying the rule in Ravn and Uhlig (2002), according to which the quarterly value of the penalty parameter λ should be adjusted by the fourth power of the frequency change (4^4 in our case), then the 400,000 value of λ suggested by Drehmann, Borio, and Tsatsaronis (2011) for financial cycles based on quarterly data leads to a λ of 1600 for annual data, which we here use.

[10] Similarly, for a specification of the determinants of the current account balance using BMA techniques, see Ca' Zorzi, Chudik and Dieppe (2012), Moral-Benito and Roehn (2016) and Desbordes, Koop and Vicard (2018).

variables that necessarily one wants to include in the model). Table A2 reports the posterior inclusion probability (PIP) of nine auxiliary regressors (GDP per capita or labour productivity; trade costs; terms of trade; government expenditure; real interest rates; NFAs; the OAD ratio;[11] private credit), where estimation is conducted via the BMA estimator introduced by Magnus, Powell, and Prüfer (2010) and the only focus regressor is country fixed effects, which need to be included in the model for the reasons discussed in Section 2. The model space is thus equal to 2^9=512 models.

In this paper, as is standard in BMA applications (e.g. Magnus, Powell, and Prüfer, 2010; Giri, Quayyum and Yin, 2019), a variable is deemed to be a relevant explanatory variable when its PIP exceeds the threshold value of 0.5 across most specifications and is thus included in the final, baseline specification of the BEER model. In this preliminary testing phase we do not impose any sign restrictions, also because the expected signs of some variables are *a priori* ambiguous. Since real interest rates, NFA and private credit display a high PIP only in maximum two cases, we exclude these variables from the BEER model.[12] Interestingly, all three variables are financial in nature, confirming the relevance of only real economic fundamentals in explaining RERs at an annual frequency since 1980.

Given their key role in open macroeconomic models, a short digression on NFAs is useful. Interestingly, this variable is excluded from the model selection even if an alternative regularization method is employed, namely the least absolute shrinkage selection operator of Tibshirani (1996). This holds true both when the penalty parameter is selected via the Bayesian Information Criterion (BIC), which assigns equal probability to all models, and when it is chosen via the Extended BIC, introduced by Chen and Chen (2008), which assumes that the more dense models are less likely. These results are available upon request. The stark insignificance of NFAs (also lagged, to take into account the fact that they are end-of-year stock measures, as in Ricci, Milesi-Ferretti and Lee, 2013) may be due to two factors. First, the steady-state relationship between NFAs and the RER is mainly expressed in the cross-section dimension and is thus difficult to detect when country fixed effects are included (Phillips et al, 2013). Second, country selection may matter. Indeed, the sample of countries considered in this paper includes fewer emerging economies than in Lane and Milesi-Ferretti's (2004) seminal article on the transfer effect, and the latter study shows how the NFA-RER relationship significantly weakens (or even turns negative) as the level of development increases. Egert, Lommatzsch and Lahrèche-Révil (2006) underscore the length of the time horizon considered in determining the sign of the link: whereas the transfer effect should hold in the very long run, in the medium term countries that experience a rapid change in their growth prospects can run current account deficits as well as having an appreciated RER, flipping the sign of the NFA-RER relationship. Since the long-run cointegrating relationship is estimated herein, this last explanation should, however, be excluded in this context.

[11] We alternately try out all four demography proxies (results available upon request). Indeed, as for the BS effect, correlated variables proxying for the same effect are not included in the same model, since the BMA methodology performs poorly in terms of model convergence when regressors are highly correlated. The OAD ratio and the aging rate have higher PIPs more frequently than the other two proxies. Given, however, that by construction the aging rate has fewer observations than the OAD ratio, we prefer to use the latter in our baseline specification.

[12] The results in Table A2 are obtained on the country sample excluding the US, in analogy with the estimation of the actual BEER model in Section 3.3; results, however, do not differ if the US is included (results available upon request).

To sum up, our baseline annual BEER model is specified as follows, one for each of the five price and cost indicators employed to deflate the dependent variable:

$$(3)\ \ rer_{i,t} = \beta_{1i}BS_{i,t} + \beta_{2i}trade_{i,t} + \beta_{3i}tot_{i,t} + \beta_{4i}gov_{i,t} + \beta_{5i}inv_{i,t} + \beta_{6\,i}oad_{i,t} + FE_i + \varepsilon_{i,t},$$

where i indicates the country, t a year in the period 1980-2017 (last year for which all relevant data are available), $BS_{i,t}$ is one of the two baseline proxies of the BS effect, $trade_{i,t}$ measures trade costs, $tot_{i,t}$ captures the terms of trade, $gov_{i,t}$ measures government consumption, $inv_{i,t}$ is the investment rate, $oad_{i,t}$ measures the OAD ratio, FE_i are fixed effects, namely country fixed effects and cross-section means of both the dependent and explanatory variables (the latter issue is discussed further on), and $\varepsilon_{i,t}$ is a random error. As mentioned earlier, all the regressors are expressed relative to the corresponding US variable (and the US is therefore dropped from the panel of countries in the estimation); for the sake of brevity, we however drop the term "relative" each time we refer to the explanatory variables.

3.3 The estimation results

In the next stage, in order to select the appropriate estimation framework, the cross-section and time-series properties of our data are investigated. The econometric details of the tests and the test results are all provided in Annex A.

As mentioned in Section 2, CSD may be an issue in the case of panel data and can raise endogeneity issues and therefore biased estimators. However, as pointed out by Pesaran (2015), only strong CSD truly poses issues in panel estimation and inference. In order to assess the presence of CSD in our panel, we thus employ Pesaran's (2015) test which tests for weak against strong CSD. Results in Table A3 show that the null is rejected for all dependent and explanatory variables, therefore suggesting the presence of strong CSD. This is unsurprising for the reasons listed in Section 2. Yet, what is surprising is that to our knowledge few panel BEER-model studies correct for this feature of the data (relevant exceptions are Bussière et al. 2010, Fidora, Giordano and Schmitz, 2018 and Giordano, 2018).

Not correctly accounting for CSD also biases panel unit root and cointegration tests. To this respect, we adopt second-generation tests which both correct for CSD and account for slope heterogeneity. In particular, we employ Pesaran's (2007) cross-sectionally augmented Im, Pesaran and Shin (2003; CIPS) test. The null hypothesis of non-stationarity of all panels of the CIPS test is tested against the alternative hypothesis that a fraction (not necessarily all) series are stationary. Results provided in Table A4 point to the null being rejected for all variables except for labour productivity, without or with the inclusion of a deterministic time trend as in Taylor (2002), Papell and Prodan (2006) and Bergin, Glick and Wu (2017). This variable is hence non-stationary for all countries, whereas the other variables are stationary for at least some economies.

Next, we perform Westerlund's (2005; 2008) group-mean cointegration test, which can be used in the presence of both stationary and non-stationary variables. It is based on the null hypothesis that the dependent and explanatory variables are not cointegrated; rejection of the null implies that these variables are cointegrated in at least some panels. Table A5 points to all differently-deflated RERs being cointegrated with the selected economic fundamentals, whichever the proxy employed for the BS effect, with the one exception of PPPs in the specification including relative labour productivity levels.

Given the tested features of our dataset, we employ the common correlated effects mean group (CCEMG) estimator – introduced by Pesaran (2006) and Kapetanios, Pesaran and Yamagata (2011), and described in Annex A – as the baseline estimator of the annual BEER model. In our view, the appropriateness of this estimation technique is due to the fact that: *(i)* it accommodates for both stationary and non-stationary cointegrated variables; *(ii)* it includes country fixed effects which, in addition to controlling for time-invariant country characteristics, are necessary due to the fact that in four out of five specifications price or cost indices are employed to construct the dependent variable; [13] *(iii)* it allows for heterogeneous slopes, which is of paramount importance given the vast country heterogeneity in our sample; *(iv)* it includes cross-section averages of the dependent and explanatory variables in order to tackle CSD. As it is a group-mean procedure, coefficients are estimated country by country and then averaged across countries.[14] The mean-group coefficients obtained from estimating equation (3) across countries are reported in Table 3, where each column refers to a differently deflated dependent variable. The top half of the table refers to estimates based on GDP per capita as a proxy of the BS effect, the bottom half on labour productivity.

The BS proxy is statistically significant only in one specification, namely that in which the RER is deflated via the PPP level (we come back to this in the robustness analysis in Section 3.4); it presents the expected positive sign. In particular, a 1 per cent rise in GDP per capita (labour productivity) of a given country relative to the US leads on average to a real appreciation relative to the US dollar of 0.46 (0.29), which is in line with the elasticities obtained in the literature (e.g. Berka and Devereux, 2013; Ricci et al., 2013; Couharde et al., 2017). Since neither BS proxy stands out as being superior, as the goodness-of-fit of the models including labour productivity are only marginally better than those including GDP per capita, both measures are retained in the baseline BEER model.

All other results also comply with the economic priors discussed earlier. In particular, an increase in the terms of trade of 1 per cent, always relative to the US, is associated with a real average appreciation relative to the US dollar of between 0.4-0.8 per cent, a result which is strongly significant across the board. A rise in trade costs and in the investment rates are also associated with a (less than proportional) real appreciation. When it is statistically significant, the long-run elasticity of government expenditure is, on average, always positive, thereby confirming the compositional bias of public spending towards the non-tradable sector, and ranges between 1.2 and 2.7, in line with estimates in De Gregorio, Giovannini and Wolf (1994). The larger magnitude we find for the ULCT-based regression is consistent with the fact that government expenditure is directed more towards the non-tradable sector and affects RERs by pushing up wages that are fully reflected in rises in the

[13] The inclusion of fixed effects removes any cross-sectional variation in the average level of the series in the panel (Fischer, 2019). Since price or cost index-based RERs do not contain any meaningful between-group information, it is quite natural in this context to employ a fixed-effects estimation, which eliminates the meaningless relative level information from the estimation process; the use of fixed-effects residuals in turn eliminates this information from the computation of the RER misalignment.

[14] The CCEMG procedure provides consistent estimates of the averages of long-run coefficients, but the latter are inefficient if homogeneity is present, as discussed in De V. Cavalcanti, Mohaddes and Raissi (2015) and Couharde et al. (2017). As we show later on, long-run slope heterogeneity is evident in our sample such that the CCEMG procedure, as opposed to a pooled procedure, is indeed warranted.

ULCT, which in contrast to the other deflators is not contaminated by developments in other cost components. Finally, also the OAD ratio is found to be positively related to RERs, with an elasticity spanning from 3.3 to 6.2.[15]

Table 3. Baseline annual BEER model estimates

A.	*Dependent variable* 1 Relative CPI	2 Relative GDP deflator	3 Relative PPP	4 Relative PPI	5 Relative ULCT
GDP per capita	0.070	0.187	0.464***	-0.150	-0.059
	0.143	0.118	0.132	0.123	0.201
Trade costs	0.320***	0.350***	0.420***	0.243***	0.040
	0.065	0.073	0.063	0.056	0.087
Terms of trade	0.526***	0.678***	0.733***	0.497***	0.694***
	0.105	0.123	0.140	0.104	0.148
Government consumption	1.275***	1.323***	1.262***	0.562	2.487***
	0.392	0.412	0.460	0.418	0.773
Investment rate	0.203	0.310	0.062	0.253	0.801***
	0.211	0.198	0.226	0.198	0.274
Old-age dependency ratio	3.666**	4.503***	3.282*	1.900	2.259
	1.676	1.715	1.992	1.524	1.845
Number of observations	*1919*	*1923*	*1923*	*1824*	*1070*
Number of countries	*54*	*54*	*54*	*52*	*37*
RMSE	*0.069*	*0.063*	*0.072*	*0.085*	*0.035*
Normalised RMSE	*0.006*	*0.006*	*0.014*	*0.007*	*0.004*

B.	*Dependent variable* 1 Relative CPI	2 Relative GDP deflator	3 Relative PPP	4 Relative PPI	5 Relative ULCT
Labour productivity	0.055	0.102	0.289**	-0.167	-0.252
	0.127	0.122	0.117	0.137	0.198
Trade costs	0.245***	0.276***	0.356***	0.165**	0.087
	0.066	0.069	0.061	0.066	0.087
Terms of trade	0.513***	0.714***	0.845***	0.357***	0.714***
	0.125	0.134	0.132	0.091	0.158
Government consumption	1.182**	1.230***	1.240***	0.635	2.655***
	0.512	0.442	0.374	0.433	0.785
Investment rate	0.541**	0.540**	0.366*	0.364*	0.428
	0.215	0.215	0.197	0.189	0.287
Old-age dependency ratio	5.204***	4.362***	3.650**	6.233***	1.699
	1.680	1.359	1.791	1.349	1.542
Number of observations	*1866*	*1866*	*1866*	*1777*	*1069*
Number of countries	*54*	*54*	*54*	*52*	*37*
RMSE	*0.055*	*0.049*	*0.064*	*0.055*	*0.036*
Normalised RMSE	*0.005*	*0.004*	*0.013*	*0.004*	*0.004*

Notes: Outlier-robust estimates obtained with a CCEMG estimator for the period 1980-2017 on the country sample excluding the US. The specifications also include country fixed effects and cross-section means, here not reported. Standard errors are reported in small font. *** p<0.01, ** p<0.05, *p<0.1. The normalised root mean squared error (RMSE) is obtained by dividing the RMSE by the (min-max) range of the dependent variable.

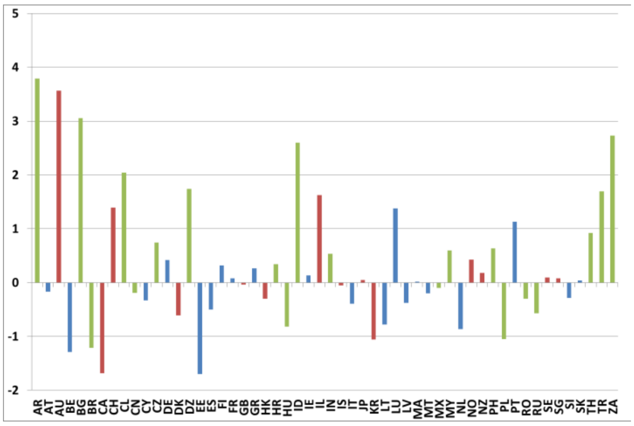
[15] These coefficients are larger than those reported in Giagheddu and Papetti (2018), but in the latter study pooled CCE estimation is conducted, thereby constraining the coefficient to be the same across all countries. This restriction is, however, not plausible, given the evidence of coefficient heterogeneity depicted in Figure 1 herein.

Figure 1 provides the country-by-country coefficient estimates for all six economic fundamentals; countries are also grouped into three categories: euro-area economies, advanced countries and emerging economies. For the sake of brevity, the figure reports estimates only for the CPI-based regression including the GDP per capita measure of the BS effect (i.e. referring to column 1, upper panel of Table 3). There is evidence of strong heterogeneity in the long-run elasticities across countries, which confirms the bias that constraining the slope coefficients to be the same, as is common under pooling procedures, would entail. The BS effect has opposite signs according to the economy considered, thereby leading to the insignificant average effect in most specifications in Table 3; it is noteworthy that for most emerging countries the relationship between the RER and relative GDP per capita is positive. Trade costs and terms of trade present positive coefficients, albeit of varying magnitude, for most economies considered. In the case of the investment rate, the positive average effect in Table 3 masks significant underlying heterogeneity, with emerging economies generally displaying a negative coefficient, plausibly pointing to the relevant role of the import content of investment, as discussed in Section 2. Finally, government consumption and the OAD ratio present very large coefficients, with opposite signs, for a bunch of countries, regardless of their level of development.

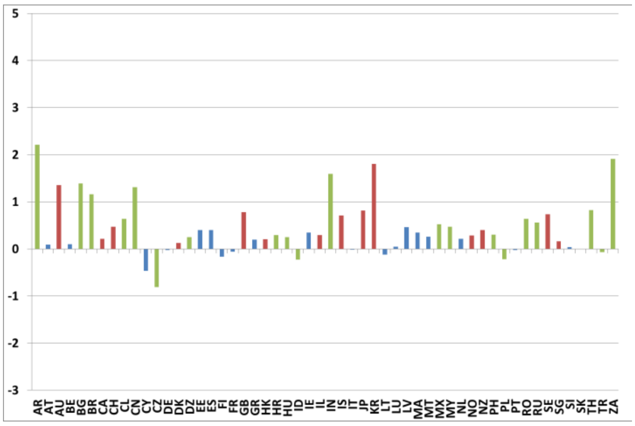
Once reassured that all relevant economic fundamentals are included in the specification and that a sound cointegrating relationship exists, goodness-of-fit statistics are in principle not meaningful in the BEER model context, since the aim of the BEER methodology is to derive a fundamental, long-term benchmark for RERs and not to maximize the fit. However, they are useful in order to exclude huge omitted variable bias and to be able to compare specifications within, and across, BEER model analyses. As in Giordano (2018), the goodness of fit of the model herein is proxied by the root mean squared error (RMSE), then normalised by dividing the standard RMSE by the (min-max) range of the dependent variable in order to guarantee comparability across specifications. The normalised RMSEs reported in Table 3 are small, pointing to satisfactory goodness of fit. They are also similar in magnitude across specifications, with only the PPP-based regression reporting a slightly higher value, possibly due to the greater measurement issues linked to PPPs than to price and cost indices. If the BEER model is satisfactorily specified, as should be the case herein, then the residual of each regression can be considered as a reasonable proxy of RER misalignment.

Figure 1. Cross-country heterogeneity of estimated coefficients

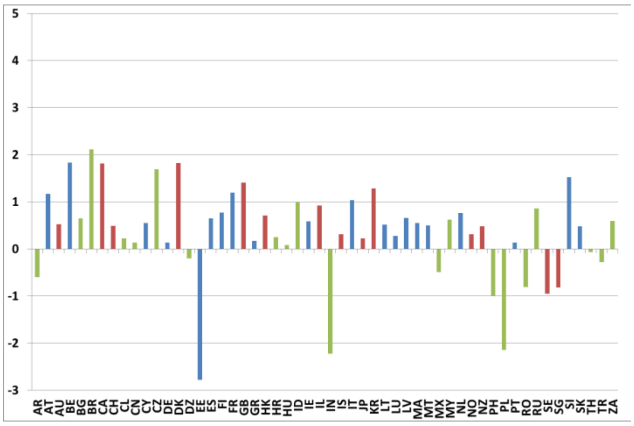
GDP per capita



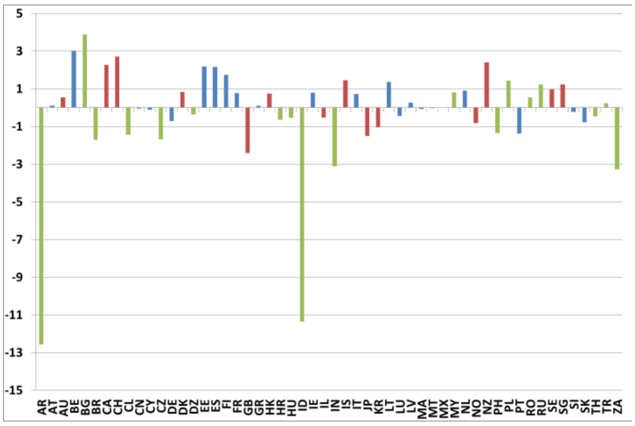
Trade costs



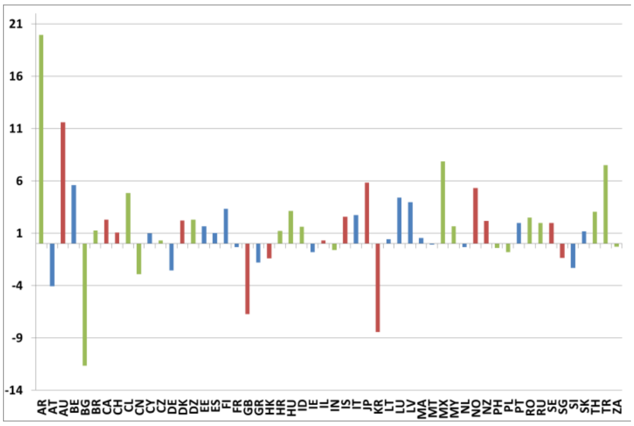
Terms of trade



Investment rate



Government consumption



Old-age dependency ratio

Notes: These country-specific CCEMG coefficient estimates are obtained from the CPI-based regression employing the GDP per capita measure as a proxy of the BS effect, shown in the first column, upper panel of Table 3. Euro-area countries are highlighted in blue, advanced economies in red and emerging economies in green; the list of countries is that in Table 1.

3.4 Robustness analysis[16]

The results in Table 3 can be shown to be robust, first, to changes in time coverage. The official nominal exchange rates used in this paper may indeed not have been the relevant benchmark

[16] All results discussed in this section are available upon request.

for transactions in some emerging economies in the early 1980s when black market exchange rates were applied (Reinhart and Rogoff, 2004). Hence, as a first robustness exercise, the sample period is restricted to the period 1985-2017. Next, since the time-span considered in this paper covers the recent double recessionary phase, which could have affected the significance and size of the link between RERs and economic fundamentals, we also estimate the BEER model on the shorter 1980-2008 period. In both cases the salience and the stability of the regressors are broadly confirmed. One noteworthy difference with respect to baseline results provided in Table 3 is that the OAD ratio is less frequently significant when focusing solely on the pre-2008 period; in other terms, the role of demography has possibly increased in most recent years, as will be later confirmed by Figure B2 in Annex B.

Baseline results are also robust to changes in the country coverage. If, for example, we exclude Argentina – which, as seen in Figure 1, reports exceptionally large coefficients, also linked to the fact that this country has displayed large exchange rate swings over time –, estimates are largely unchanged. Further changes in the country sample are discussed below.

In order to specifically address the concern in Hlouskova and Osbat (2009) relative to a possible bias deriving from a systematic misalignment in the numeraire country, discussed in Section 3.1, we include a time trend in specification (3). The trend is never significant and indeed its inclusion does not alter the baseline coefficient estimates. Moreover, we also exclude those countries that were *de facto* pegged to the US dollar over a significant part of the estimation period; indeed, in these countries RER developments were muted against potentially large movements in relative economic fundamentals. In order to identify the countries with a pegged currency *vis-à-vis* the US dollar in our country sample, we adopt Shambaugh's (2004) classification, according to which a country is considered a "pegger" if its exchange rate fluctuates within a +/-2 percent band against a base currency (i.e the currency with historical importance for the country, the nearby dominant economy to which other currencies are pegged, or the dollar as a default). As a result the countries excluded in this robustness check are China, Hong Kong and Malaysia. Baseline results are again broadly confirmed, adding to the studies that find that the choice of the numeraire does not significantly affect BEER-model estimation.

Given its importance, we also investigate the BS effect in more detail. First, we better explore its weak significance in Table 3. We exclude the possibility of strong collinearity with the investment rate, in that, when omitting the latter variable from the BEER model, the BS effect is still found to be insignificant in seven out of ten specifications (the two PPP-based specifications and one GDP deflator-based specification). From a theoretical standpoint, it could be argued that our result is in line with the predictions of Benigno and Thoenissen's (2003) model, which claims that when productivity increases in the traded goods sector, an endogenous depreciation of the terms of trade (held constant in the original BS model) offsets the appreciation of the relative price of non-traded goods, implying that supply-side improvements to productivity result in a depreciation of the RER, in contrast with the BS proposition. However, this is not the case herein as terms of trade are also included in the BEER model: hence, conditional on terms of trade, the traditional BS mechanism should apply. (Incidentally, when terms of trade are dropped, the BS effect is still insignificant in seven out of ten specifications). More convincingly, from an empirical standpoint, the existing literature finds greater support for the BS hypothesis in cross-section than in time series. Indeed, if we lower the frequency of the observations by constructing five-year non-overlapping averages of the two BS proxies and re-estimate the BEER model, similarly to Berka and Steenkamp (2018), the BS effect comes out as significant in the three broad-sample deflated regressions (CPI, GDP deflator and PPP), but still remains insignificant in the PPI and ULCT specifications, confirming the relevance of the cross-section, as opposed to the time-series, dimension in assessing the BS effect.

Next, as in Schnatz, Visjelaar and Osbat (2003), Bénassy-Quéré, Béreau and Mignon (2009) and Fidora, Giordano and Schmitz (2018), we investigate alternative proxies and representations of the BS effect. First, we consider an indirect proxy of relative productivity in the tradable sector, namely the CPI-to-PPI ratio. The intuition is that, unlike the CPI which includes, for instance, services and housing, the PPI broadly covers only tradable goods and therefore the CPI-to-PPI ratio proxies the non-tradable vs. tradable price ratio. Relative to the two direct indicators of the BS effect employed thus far, this proxy has the advantage of considering relative sectorial developments. However, in addition to the fact that this ratio is an imperfect measure of the non-tradable vs. tradable price ratio (Engel, 1999; Chinn, 2006), there are other factors that might affect relative price developments and that are not related to productivity, such as changes in value-added taxes that affect CPI but not PPI, biasing the ratio. We find that this proxy is always statistically significant and with the correct positive sign in the PPP, GDP deflator and ULCT-based regressions,[17] but given that PPIs are not available for all countries in our sample and given that this proxy limits the choice of the dependent variable, we prefer not to include this third measure in our baseline model. Second, we employ gross national product (GNP) per capita as opposed to GDP per capita, in order to control for the fact that some countries in our sample may be associable to tax havens. GNP series are, however, only available since 1990. The BS effect is statistically significant in two out of five specifications (PPP and GDP deflator-based), with a comparable magnitude to the GDP per capita measure reported in Table 3, and the coefficients of the other regressors are also in line with those presented in Table 3. However, given limited data availability for GNP per capita, this variable could not be employed as a baseline proxy.[18] Third, we test the accuracy of a linear representation of the BS effect. Indeed, Kravis, Heston and Summers (1978) and, more recently, Kessler and Subramanian (2014) and Hassan (2016) uncover non-linearities in the relationship between PPP-deflated RERs and relative GDP per capita levels over long time-spans: they find that the BS effect holds only for middle- and high-income countries, whereas the relationship is negative for low-income countries. In particular, Hassan (2016) explains this result by arguing that increases in productivity in agriculture lead to decreases in the relative price of agriculture and, in turn, of the aggregate price level in low-income countries, as their share of agriculture in total labour is high. Only above a certain income threshold, productivity in manufacturing relative to services becomes the main driver of the aggregate price level and the standard BS effect can thus be detected. We therefore augment specification (3) with second-order terms of the two alternative baseline BS measures. The quadratic term however is significant only in one specification out of ten, such that we conclude that a linear formulation is appropriate to proxy for the BS effect in our sample, possibly due to the fact that the latter does not include low-income countries. This finding was also highlighted in Fidora, Giordano and Schmitz (2018).

[17] In the CPI-based regression it is positive and large, whereas in the PPI-based regression it is negative, largely by construction.

[18] Alternatively, and more roughly, when countries such as Luxembourg or Ireland are dropped from the estimation sample, results are largely unvaried to those reported in Table 3.

Finally, we check for the presence of further omitted variables to those discussed earlier. In particular, we augment equation (3) with a measure of railway density, namely the ratio of railroad coverage to land density. Indeed Du, Wei and Xie (2013) find that a (relative) improvement in transportation infrastructure, which leads to a drop in domestic prices via a competition effect, may be nearly as important in magnitude as the BS effect in shaping the RER. Although the BMA analysis does not point to the need to include railroad density amongst the economic fundamentals in the BEER model (results available upon request), we also double-check by augmenting specification (3) with this variable to assess its statistical significance conditional on our selected set of regressors. This additional variable is, however, never statistically significant in any of the ten specifications, possibly due to the fact that our country sample contains relatively too few emerging economies; only in the latter countries transportation infrastructure progress is presumably of appreciable magnitude to be able to exert a significant impact on the RER.

To sum up, the robustness and sensitivity analysis conducted in this section point to the soundness of the baseline BEER model, from which RER equilibrium values can therefore be derived, as discussed in the following section.

3.5 Deriving equilibrium values

The equilibrium RER values for all 55 countries, including the US, are obtained as fitted values based on the average coefficients in Table 3.[19] The use of average coefficients is necessary to derive a consistent and symmetric set of equilibrium exchange rates (Bénassy-Quéré, Lahrèche-Révil and Mignon, 2008); country-specific coefficients would in fact violate bilateral consistency, as a given change in a (relative) economic fundamental of a given country against the numeraire country would potentially not have a symmetric impact on the bilateral exchange rate between the two countries.

Two different concepts of equilibrium exchange rates can be derived, depending on whether the explanatory variables are set at their observed values or at their "equilibrium" level (e.g. MacDonald, 2000). In the latter case, one corrects for the possible misalignments of the economic fundamentals themselves. In turn, two main strategies may be used to obtain long-term equilibrium values for the fundamentals. On the one hand, one can draw on theory and rely on a structural model for each economic fundamental. For example, Bénassy-Queré, Béreau, and Mignon (2008) use a model to estimate the equilibrium value of the NFA position, which is then employed in their BEER model. However, this approach is very challenging when manifold fundamentals are included in the BEER model, as is the case in this paper. On the other hand, and more commonly, atheoretical filtering techniques can be employed.[20] Starting with Clark and MacDonald's (1998) seminal study, the Hodrick-Prescott (HP) filter, for example, has been used by Nouira and Sekkat (2012), Schröder

[19] The US's equilibrium RER is computed on the basis of the mean-group constant and of the cross-section means.

[20] These decomposition techniques are not, however, devoid of caveats. In emerging economies some economic fundamentals, such as government consumption, may have been persistent but unsustainable (e.g. Baffes, O'Connell and Elbadawi, 1999). In these cases, unsustainable values would be passed through to the trend component. A means of limiting this possible bias is to exploit as wide a sample period as possible, as is done herein.

(2013) and Comunale (2017). In this paper too the baseline equilibrium RER values are obtained on the basis of HP-filtered explanatory variables.[21]

The equilibrium REER is then computed by weighting the equilibrium RERs with three-year time-varying trade weights sourced from the ECB (generally, *vis-à-vis* 54 countries; see Table 2 for details). The REER misalignment is then gauged relative to the predictions of the BEER model: the difference, in percentage points, between the actual and the equilibrium REER is labelled as the REER misalignment, as follows:

$$(4)\ mis_{i,t} = reer_{i,t} - reer_{i,t}{}^{*}$$

where the asterisk denotes the equilibrium value. Given how the REER is defined, when the misalignment is negative (positive), and therefore the actual REER is more depreciated (appreciated) than the equilibrium REER, it implies an undervaluation (overvaluation) of the actual REER.

For each country/year we generally obtain ten REER equilibrium values, according to each different deflator and one of two alternative BS proxies, and thus ten REER misalignment estimates (recall that ULCT-based estimates are only available since 1995 and for only 38 countries; PPI-based estimates are only available for 53 countries). The same procedure applies to euro-area countries' HCIs. Table 4 provides some summary statistics of the estimated annual misalignments, based only on the GDP per capita BS proxy for the sake of brevity. On average since 1980, these indicators have been undervalued in the range of 9 (PPP-based estimates) to 38 (ULCT-deflated estimates) per cent, yet with large variation across countries and time; undervaluations have been more frequent than overvaluations, regardless of the deflator considered.

Table 4. Descriptive statistics of REER/HCI misalignments

	Mean	St. dev.	Min		Max		N. obs	N. positive	N. negative
CPI-based REER/HCI	-36.5	77.7	-388.7	(HU 1991)	172.5	(AR 1986)	2004	878	1126
GDP deflator-based REER/HCI	-36.2	75.8	-385.7	(HU 1991)	173.8	(AR 1986)	2001	812	1189
PPP-based REER/HCI	-9.3	28.8	-289.7	(BG 1991)	173.6	(AR 1986)	1988	844	1144
PPI-based REER/HCI	-29.6	72.2	-362.7	(HU 1995)	233.6	(IL 1983)	1888	805	1083
ULCT-based REER/HCI	-38.3	73.4	-360.7	(HU 1995)	45.6	(BG 2017)	874	263	611

Source: author's estimates, based on the annual model described herein, employing the GDP per capita measure as a proxy of the BS effect.

In Figure B1 in Annex B we compare both our five-deflator average and our CPI-based HCI misalignments of the four main euro-area economies with the CPI-based measures produced by CEPII. To our knowledge, CEPII is the only other institution that publishes long-run estimates for many countries according to an annual BEER model, the "EQCHANGE model" whose latest version is described in Grekou (2018a), and which is similar in spirit to that developed in this paper.[22]

[21] The λ smoothing parameter is selected according to the Ravn-Uhlig rule discussed in footnote 9 and is equal to the conventional value of 100 (see, for example, Schröder, 2013).

[22] In a nutshell, the latest EQCHANGE model is based on 143 countries at an annual frequency over the period 1973-2017. The dependent variable is the CPI-based REER and the economic fundamentals included are: relative GDP per capita, NFAs, terms of trade, government spending and trade openness. Estimation is conducted via the pooled mean group estimator. See Giordano (2018) for a systematic comparison of the CEPII, IMF and Bank of Italy REER misalignment estimation methods and results, as well as De Nardis (2018) for a discussion of the resulting estimates for selected euro-area countries.

Especially when comparing CPI-based estimates only, results are qualitatively comparable (and strikingly similar in the case of Germany), in particular in the most recent years.[23]

4. The quarterly BEER model

In this section we briefly describe a quarterly BEER model that is comparable with the annual model discussed in the previous section. It is a variation of that depicted in Giordano (2018), which is currently employed at the Bank of Italy for monitoring and analysis purposes. In particular, the main changes consist in: a) estimating the model on the 55 countries listed in Table 1, as opposed to 57;[24] and b) adopting the US, as opposed to the euro area, as the numeraire country. All other features of the original model are maintained, namely the length of the estimation period (1999Q1-2017Q4) – which is anyhow long enough to assess long-run relationships (Engel and Zhu, 2019) –, the selection of the data sources (discussed in Giordano, 2018, to which we refer), the employment of five alternative price/cost indicators to deflate the RER and the choice of the estimation technique (CCEMG).

These modelling choices therefore guarantee full comparability between the two models discussed in this paper, which thereby differ only by the data frequency and its implications, namely the shorter length of the estimation period (as of 1999, as opposed to 1980) and the different set of relevant economic fundamentals in explaining the bilateral RERs.[25] Fidora, Giordano and Schmitz (2018), the first version of the model outlined in Giordano (2018), thoroughly discusses the selection method of the significant quarterly explanatory variables, which leads to including (short-term) interest-rate differentials, but omitting the relative investment rate and demographic variables. Although the mentioned study adopted a general-to-specific approach to select the relevant fundamentals, the same set of explanatory variables would have been chosen according to BMA (results are available upon request). In particular, the fact that the real interest rate is found to be significant in tracking the bilateral RER at the quarterly frequency is consistent with the conclusions generally obtained in the literature suggesting that RER movements are related to real interest-rate differentials over medium-term rather than long-term horizons (e.g. Campbell and Clarida, 1987; Meese and Rogoff, 1988; MacDonald, 1998; Bénassy-Quéré, Béreau and Mignon, 2009). Conversely, slow-moving factors, such as the OAD ratio, are insignificant. Moreover, at the quarterly frequency only the GDP per capita proxy of the BS effect is statistically significant, in several specifications; this may be due to the fact that labour productivity is more cyclical at the quarterly frequency than GDP per capita, due to episodes of labour hoarding or shedding.

The specification of the quarterly model is the following, for each of the five price and cost indicators employed to deflate the dependent variable, the bilateral RER *vis-à-vis* the US dollar:

[23] One exception is Italy in the most recent years, for which CEPII estimates point to an overvaluation, albeit significantly smaller than that recorded in the 2000s.

[24] In particular, Venezuela has been dropped owing to the recent suspension of the publication of its official national accounts and price series, as well as Taiwan, owing to the lack of data availability for some variables employed in this paper.

[25] Moreover, since the phi-ness indicator of trade costs discussed in Section 3.2 and employed in the annual model is not available at a quarterly frequency, it is here replaced by the more standard measure of openness to trade (which, however, has been shown to be highly correlated with the former measure in Figure A1).

$$(5)\quad rer_{i,t} = \beta_{1i}GDPpercap_{i,t} + \beta_{2i}tradeopen_{i,t} + \beta_{3i}tot_{i,t} + \beta_{4i}gov_{i,t} + \beta_{5i}intrate_{i,t} + FE_i + \varepsilon_{i,t},$$

where i indicates the country, t a quarter in the period 1999Q1-2017Q4, $GDPpercap_{i,t}$ is GDP per capita, $tradeopen_{i,t}$ measures trade openness, $intrate_{i,t}$ is the real short-term interest rate, $tot_{i,t}$ and $gov_{i,t}$ are defined as before, FE_i are fixed effects, namely country fixed effects and cross-section means of both the dependent and explanatory variables, and $\varepsilon_{i,t}$ is a random error. All the regressors are expressed relative to the corresponding US variable, and the US is dropped from the sample.

The coefficients obtained from estimating equation (5) across countries via CCEMG are reported in Table 5, where each column refers to a differently deflated dependent variable. Results are qualitatively very similar to those reported for the annual model in Table 3. The BS effect is found to be more frequently statistically significant than in the annual version, yet presents a smaller coefficient. Openness to trade, similarly to the phi-ness indicator in Table 3, is highly significant and presents the correct sign (negative, as it is an inverse proxy of trade costs). The coefficients attached to terms of trade and government consumption both display the expected positive signs, with the latter confirming its large magnitude especially in the ULCT-based regression. Finally, real interest rates are found to be significantly and positively associated with RERs.

The goodness of fit of the quarterly specifications is satisfactory, with the PPP-based regression displaying a marginally higher normalised RMSE than the others. Overall, the normalised RMSEs are lower in the quarterly, as opposed to the annual, model.

Table 5. Baseline quarterly BEER model estimates

	Dependent variable				
	1	*2*	*3*	*4*	*5*
	Relative CPI	Relative GDP deflator	Relative PPP	Relative PPI	Relative ULCT
GDP per capita	0.128	0.310**	0.266*	0.0282	0.378***
	(0.141)	(0.156)	(0.152)	(0.130)	(0.136)
Openness to trade	-0.457***	-0.545***	-0.490***	-0.377***	-0.330***
	(0.0871)	(0.0928)	(0.0926)	(0.0744)	(0.0900)
Terms of trade	0.350***	0.524***	0.628***	0.255**	0.252**
	(0.122)	(0.130)	(0.127)	(0.101)	(0.122)
Government consumption	1.122**	0.994*	1.428***	0.855*	1.722***
	(0.532)	(0.563)	(0.520)	(0.498)	(0.362)
Real interest rate	0.005***	0.006***	0.007***	0.004***	0.003**
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Number of observations	*4104*	*4104*	*4104*	*3952*	*2812*
Number of countries	*54*	*54*	*54*	*52*	*37*
RMSE	*0.027*	*0.027*	*0.027*	*0.026*	*0.023*
Normalised RMSE	*0.003*	*0.003*	*0.004*	*0.002*	*0.003*

Notes: Outlier-robust estimates obtained with a CCEMG estimator for the period 1999Q1-2017Q4 on the country sample excluding the US. The specifications also include country fixed effects and cross-section means, here not reported. Standard errors are reported in small font. *** p<0.01, ** p<0.05, *p<0.1. The normalised RMSE is obtained by dividing the RMSE by the (min-max) range of the dependent variable.

At this point quarterly misalignment estimates can be produced according to the procedure discussed in Section 3.5, with the difference that, since only one BS proxy is significant in the quarterly model, for each country-quarter five, as opposed to ten, estimates are available. Moreover, the economic fundamentals in the quarterly model are not filtered, since applying filtering techniques to the quarterly data makes no significant difference to the resulting misalignments (results available upon request), as already found in Fidora, Giordano and Schmitz (2018). In the next section we compare our estimated annual and quarterly misalignments.

5. HCI and REER misalignments: a comparison across different data frequencies

5.1 HCI misalignments of the four main euro-area economies

In order to gain some insight into the differences in estimates stemming from the two models described thus far, in this section we focus on the four largest euro-area economies.

It is first interesting to analyse the long-run developments of HCI misalignments based on the annual model (Fig. 2). In the early 1980s the uncertainty surrounding the annual misalignment estimates of all four countries was historically high – possibly due to the worse quality of PPPs, but also of the PPIs, in the initial years of the sample period – and the size, but even the sign, of the disequilibria are unclear.[26] As of the mid-1980s, all four economies recorded rising overvaluations (with the exception of France, where the overvaluation was broadly flat), peaking on average at around 20 per cent in both Italy and Germany in the early Nineties. The 1992 currency crisis led to an abrupt absorption of the large misalignments in Italy and Spain in that same year; in Germany, which was also concurrently undergoing the immediate effects of political unification, the downward correction was more gradual and spread out over the whole decade.

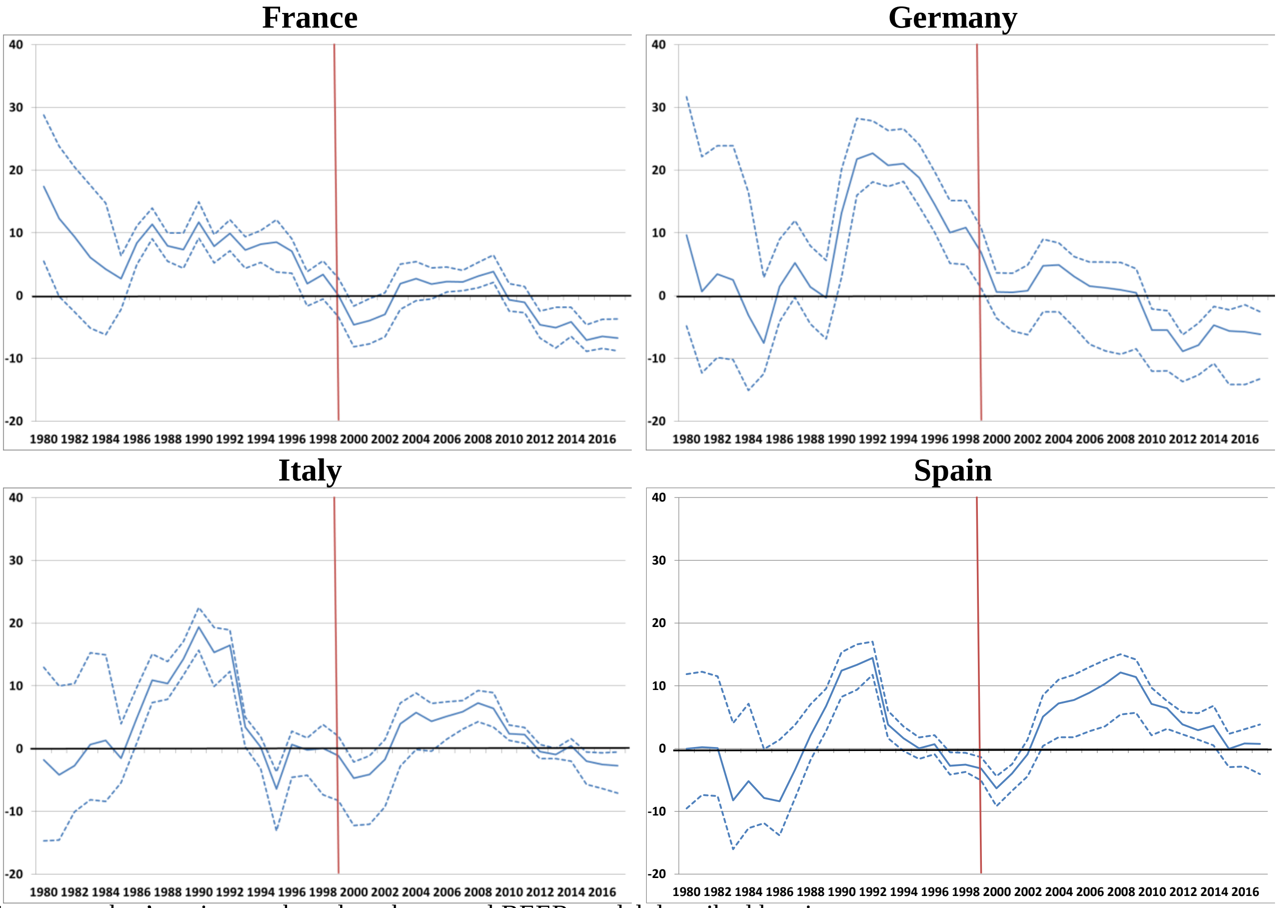
After the inception of the Economic and Monetary Union (EMU) and until 2017, the development of misalignments marked a clear inverted U-shape in France, Italy and Spain, whereas it followed a negative trend in Germany. In particular, until the outbreak of the recent global financial crisis, average misalignments turned from negative to positive in France, Italy and Spain, reaching their maximum levels on average at over 12 and 7 per cent in 2008 in Spain and Italy, respectively.[27] Conversely, in the same years Germany's HCIs were on average broadly in line with economic fundamentals.

[26] As discussed more thoroughly in Giordano (2018), the IMF-WEO PPP rate, which is used in this paper as one of two PPP sources, is obtained by retropolating the 2011 PPP rate with the GDP deflator growth rate; the further the period relative to the benchmark year, the less precise is the measured PPP level. Concerning the PPI instead, sourced from the Bank of Italy, as documented in Felettigh and Giordano (2018), the earlier the years in the sample the more heterogeneous are the sources and the price aggregate employed. The 1980s are also very particular years (e.g. Turner, 1991). The stringent capital restrictions, introduced in most advanced economies under the Bretton Woods Agreement and still present in the 1970s, were loosened in the 1980s. At the same, innovation in both communication systems and financial transactions contributed to a boom in aggregate capital flows, especially amongst advanced countries. A rise in the volatility of these flows was also registered relative to the previous decade; this could be linked to the uncertainty in the REER misalignment measures in Figure 2. Indeed, in Turner (1991; p. 102) it is argued that in the 1980s "prices in virtually all markets – equities, bonds and foreign exchange – have become much more volatile, And disturbances in one market have been more quickly transmitted to other markets, often irrespective of the underlying fundamentals".

[27] Whereas Spain's two episodes of large overvaluation (early Nineties; build-up to the global financial crisis) were broadly comparable in magnitude, in Italy the second episode was less pronounced than the first.

Figure 2. HCI misalignments of the main euro-area countries (1980-2017) according to the annual model

(percentage points)



Source: author's estimates, based on the annual BEER model described herein.

Notes: The full lines represent the mean of the ten estimated HCI misalignments across five price/cost indices and two alternative BS proxies for each reference country; the dashed lines represent the minimum and the maximum. A positive (negative) misalignment would require a depreciation (an appreciation) of the actual HCI to unwind, provided the equilibrium rate does not change.

Since 2010, all HCI measures for Germany have pointed unanimously to a substantial undervaluation, which peaked on average in 2012 at nearly 9 per cent. France too recorded a significant undervaluation in the 2010-2017 period, whereas Italy and Spain's average HCI misalignments went from both positive to slightly negative and to broadly zero, respectively.

On the whole, average HCI misalignments in the four main euro-area countries were significantly smaller after 1999 than in the preceding period, dropping from a median misalignment across the four economies of nearly 12 per cent in the years 1980-1998 to around 5 per cent in the period 1999-2017. This result also holds when considering the twelve countries that had adopted the single currency by 2001.[28] Indeed, the adoption of the euro may have lowered HCI misalignments in the member countries due to various factors, amongst which: the structural measures undertaken in order to meet the Maastricht criteria in the Nineties, which lowered inflation rates, and differentials, as a result (Berka and Devereux, 2013); the enhanced intra-euro area trade stemming from the

[28] Results also hold when Luxembourg is dropped, as in Fidora, Giordano and Schmitz (2018).

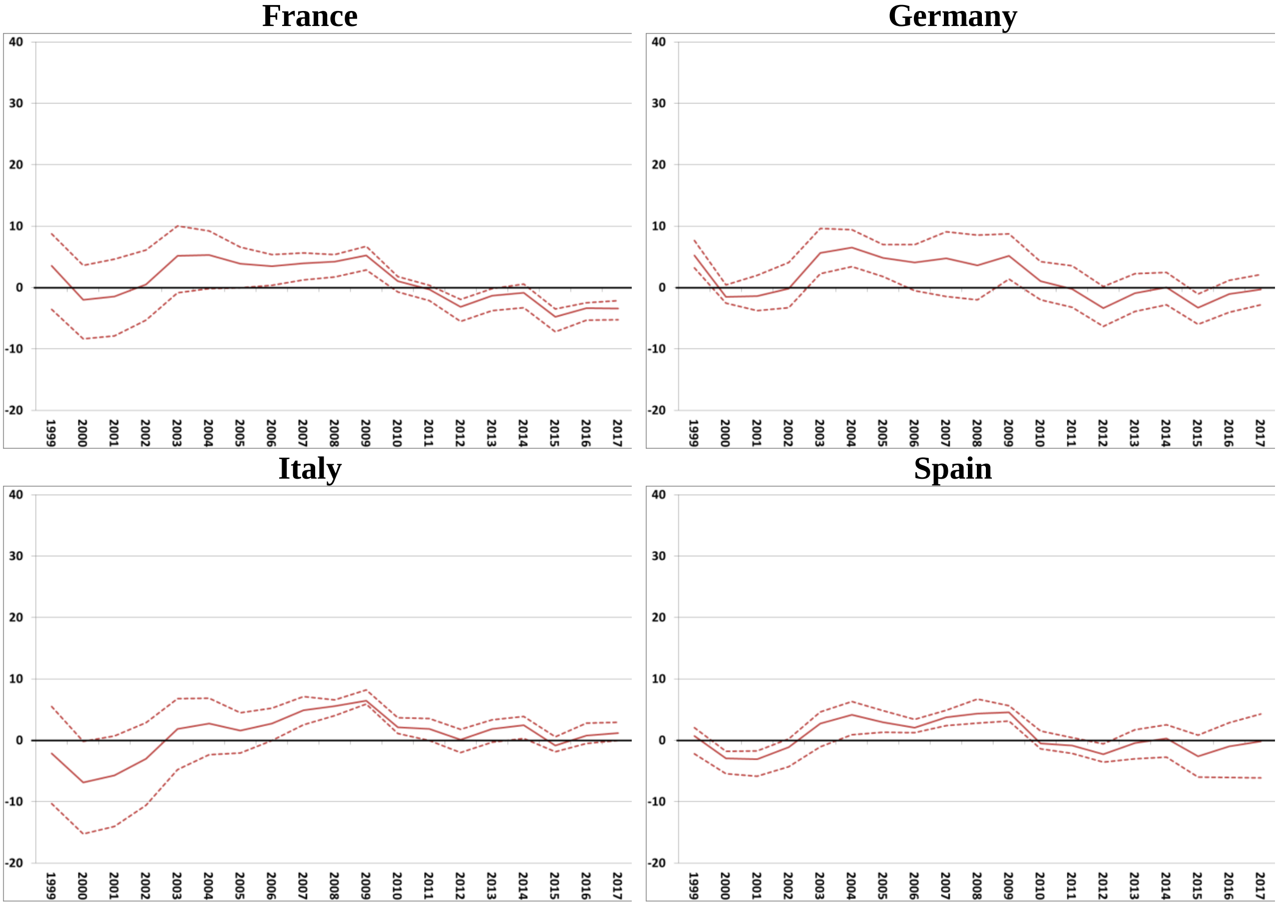
monetary union, which could have accelerated price convergence across euro-area countries; the elimination of the nominal exchange rate channel, a potential source of volatility stemming from financial-market turbulence, which can in turn exacerbate REER misalignments (Bergin, Glick and Wu, 2017).

These estimates therefore corroborate the cross-sectional findings in Fidora, Giordano and Schmitz (2018), according to which median misalignments in the 1999-2016 period were significantly lower in the founding members of the euro area than in non-euro area advanced and emerging economies. They are also consistent with the recent findings in Engel and Zhu (2019), which point to a lesser disconnect between the RER and economic fundamentals under rigidly fixed nominal exchange rates than under floating rates. They instead run counter to results in Coudert, Couharde and Mignon (2013), which point to an increase in HCI misalignments on average for euro-area countries since the inception of the EMU. The latter estimates are, however, amongst other things, based on a very sketchy BEER model, inclusive of only two economic fundamentals (i.e. a BS effect proxy and NFAs), with a potentially large omitted variable bias, and on HCIs computed *vis-à-vis* solely 27 trading partners (against 54 in this paper, and 56 in Fidora, Giordano and Schmitz, 2018). Moreover, increasing HCI misalignments are not found for all euro-area countries in Coudert, Couharde and Mignon (2013): Germany and Italy, for example, display decreasing misalignments after 1999 even in that study.

Figure 3 reports annualised misalignment estimates based on the quarterly BEER model, for the 1999-2017 period. Medium-term developments are similar to those obtained according to the annual model: the generally rising overvaluation until the eruption of the global financial crisis and the subsequent downward correction which in the case of Germany and, especially, France, was particularly deep, leading to a significant average undervaluation. Point estimates for the more recent years differ slightly across the two models: Germany’s current undervaluation is less pronounced in the quarterly than in the annual model (as is the post-2004 downward trend in misalignment), and Italy’s HCI appears to be marginally overvalued on average in 2017, against a mild undervaluation according to the annual model.

Figure 3. HCI misalignments of the main euro-area countries (1999-2017) according to the quarterly model

(annual averages of quarterly estimates; percentage points)



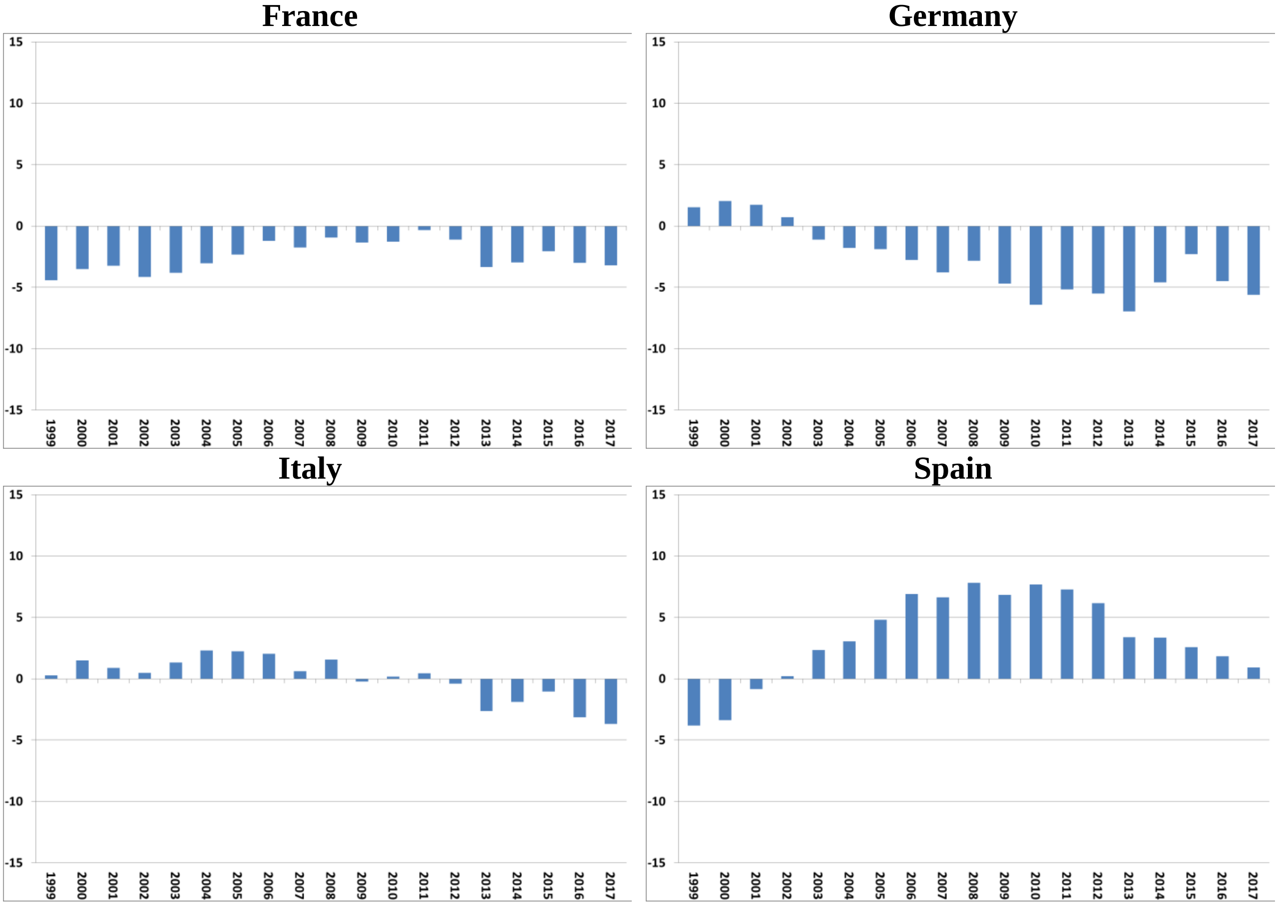
Source: author's estimates, based on the quarterly BEER model described herein.

Notes: The full lines represent the mean of the five estimated HCI misalignments across five price/cost indices for each reference country; the dashed lines represent the minimum and the maximum. A positive (negative) misalignment would require a depreciation (an appreciation) of the actual HCI to unwind, provided the equilibrium rate does not change.

Figure 4 helps further assessing the differences between the two models, by providing the percentage-point gap between the two sets of mean estimates. First, the differences are not systematically signed, since they vary across countries and years. Second, if one considers the high statistical uncertainty surrounding BEER model estimates (see, for example, the discussion in Cheung, Chinn and Fujii, 2010), these gaps are generally contained, in particular for France and Italy, whereas they are more pronounced for Spain, but only in 2006-2012, and, to a lesser extent, in Germany since 2010.

Figure 4. Mean HCI misalignments of the main euro-area countries (1999-2017) according to the annual and quarterly models: a comparison

(percentage-point differences between annual and annually-averaged quarterly mean estimates)



Source: author's estimates, based on the annual and quarterly models described herein.

Notes: The bars represent the percentage-point difference between the annual-model and the annualised quarterly-model mean estimates (i.e. the full lines in Figure 2 and Figure 3, respectively).

Although it is far from trivial to gauge the causes of these more marked discrepancies, Figure B2 in Annex B sheds further light into this issue by providing the "pseudo-contributions" of each economic fundamental to the HCI equilibrium paths of the four main euro-area countries since 1999. We employ the term "pseudo-contributions" owing to the fact that computing precise contributions in our annual BEER model is analytically cumbersome in that, as mentioned in Section 3, CCEMG estimation also includes cross-section means of all the economic fundamentals (which therefore mire the calculation of the contribution of each fundamental), as well as of the dependent variable, in order to control for CSD.[29] In Figure B2 for each country we bunch together the intercept and these CSD means (where the latter can be considered as "exogenous" common shocks which affect the domestic HCI), whereas we provide the pseudo-contribution of each "domestic" economic fundamental separately.[30] This breakdown, although imprecise, is not totally uninteresting since it quantifies the

[29] In order to quantify the contribution of economic fundamentals in their annual BEER model, Giagheddu and Papetti (2018), for example, revert to a Dynamic Ordinary Least Squares (DOLS) specification which, however, is biased in the presence of CSD (see, for example, Giordano, 2018 for a comparison of a CCEMG vs DOLS estimation of a quarterly BEER model). This option is here not considered, given the proven presence of CSD.

[30] For example, the pseudo-contribution of the BS proxy is $\beta 1 BSi,t$, where $\beta 1$ is the group-mean coefficient attached to the BS proxy in that, as explained in Section 3.5, it is this coefficient that is employed to compute the equilibrium values.

role of the latter variables, the only variables on which domestic policy-makers can intervene. For the sake of brevity, we solely consider results referring to the CPI-deflated HCI equilibrium rate, based on the specification including GDP per capita as the proxy of the BS (Table 3, upper panel, col. 1).

In general, in the four economies, in addition to the constant and CSD variables, trade costs and the OAD ratio are the most relevant contributors to the annual equilibrium HCI. Interestingly, the contribution of these two variables, displaying opposite signs, have increased in absolute value over time, with the only exception of that of the OAD ratio in Spain, which was broadly stable. According to the country and the year, government consumption and terms of trade also played a significant role. The real interest rate, which is included only in the quarterly model, instead has a negligible contribution to the equilibrium HCI, given its tiny coefficient reported in Table 5, and therefore contributes little to explaining the differences.

It is noteworthy that in the second half of the 2000s Spain's investment rate (relative to the US) crashed, contributing to dampen the country's HCI equilibrium value more according to the annual, as opposed to the quarterly, model, and thereby revealing a larger overvaluation, all other things equal. Instead, Germany's OAD ratio decreased relative to that of the US in the 2010-2017 period, passing from a positive contribution of 0.43 to 0.34 points. This development significantly pushed the country's annual equilibrium level down (against broadly flat dynamics of the quarterly-model-based equilibria), and reduced the size of its HCI undervaluation relative to the quarterly model.

In the next section we turn to one way of empirically assessing the informative content of the two sets of misalignments put forward herein.

5.2 Misalignments as predictors of REER developments

Thus far, we have estimated REER misalignments in a (panel) cointegration framework, labelling these imbalances as deviations of actual REERs from their long-run equilibria; in the presence of a misalignment, this framework thereby assumes a future correction of the actual REER to its equilibrium value (which may however evolve in the meantime). In this section we assess whether annual and quarterly misalignment estimates have a different explanatory power for subsequent observed REER movements in the whole sample of 55 countries. Following Abiad, Kannan and Lee (2009), this exercise helps appraising the usefulness of REER misalignment estimation according to either model in terms of predicting expected future REER changes, and therefore each model's potential contribution to economic analysis and policy-making.

In detail, we estimate the reactivity of actual REERs to past misalignments. Since deviations from equilibrium levels can also be narrowed down by suitable changes in economic fundamentals, reducing the necessary adjustments in the REERs, we also include the change in fundamentals as a control variable.[31] Because of the limited number of years underlying the quarterly model, we only perform this exercise in-sample. We thus estimate the following regression in a standard fixed-effects panel regression setting:

[31] For similar exercises, see also Salto and Turrini (2010), Yeşin (2016), Adler and Grisse (2017), Fidora, Giordano and Schmitz (2018).

(6) $\Delta reer_{i,t/(t-5y/20q)} = \alpha_i + \beta_1\, mis_{i,t-5y/20q} + \gamma_1\, \Delta fun_{i,t/(t-5y/20q)} + \varepsilon_{i,t}$

where α_i are country fixed effects, the coefficient β_1 measures the sensitivity of REER changes to past (i.e. lagged by five years, y or 20 quarters, q) misalignments, $mis_{i,t}$ is defined in equation (4), γ_1 estimates the sensitivity of changes in the actual REERs to contemporaneous changes in fundamentals (i.e. in the equilibrium path) and $\varepsilon_{i,t}$ is a random error. Five-year windows are considered in order to allow for sufficient time to observe real adjustments take place. In order not to sacrifice too many observations, which would hinder the estimation of equation (6) based on quarterly data, we adopt overlapping windows, as in Abiad, Kannan and Lee (2009).

Estimation results are provided in Table 6 for each of the REER deflators and for either of the BS proxies. The upper panel refers to results based on annual estimates, obtained via estimation over the entire 1980-2017 period; the middle panel refers to results based on annual estimates, obtained via estimation over the restricted 1999-2017 period, so as to be comparable to the quarterly model's time span; the lower panel refers to results based on quarterly estimates, obtained via estimation over the 1999-2017 period. The robust standard errors reported in Table 6 address the serial correlation stemming from overlapping observations (see Hansen and Hodrick, 1980 for a discussion of this issue).

The following results stand out. First, β_1 is statistically significant and negative across all specifications. This finding is reassuring in that it points to the estimated REER misalignments having predictive power over future movements in actual REERs; in other terms, actual REERs tend to converge towards their (lagged) estimated equilibrium values. In particular, more overvalued currencies tend to experience subsequent larger real depreciations; conversely, more undervalued currencies tend to record larger real appreciations. Interestingly, β_1 is larger in the case of the quarterly-model estimates, and this result is not due to the shorter estimation window since it also generally holds when the annual-model estimation period is restricted to the 1999-2017 period. Wald test results available on request point to β_1 being significantly different from -1 (which entails full correction over the selected time horizon) in all annually- and quarterly-specified regressions, implying that five years are not sufficient for the misalignments to fully recede. This result is therefore consistent with the lower bound of the range of estimated macro half-lives (three to five years) reported in Rogoff's (1996) seminal article.

Second, the coefficient γ_1 also displays the expected positive, statistically significant sign across all specifications, signalling that an appreciation of the equilibrium REER due to changes in economic fundamentals is associated with an appreciation of the actual REER. Third and most importantly, measures of misalignment based on the quarterly model appear to have a significantly higher explanatory power, in terms of larger adjusted R^2, of subsequent REER developments than misalignment estimates stemming from the annual model. This result could be due to the higher number of observations, but also to the high-frequency intrinsic nature of REERs.

Finally, the described results (available upon request) are also robust to three-year windows. In this case, the magnitude of β_1 is smaller across the board, as expected, since the shorter the period considered, the smaller is the observed correction. Moreover, with three-year windows, misalignments – however measured – explain a smaller share of actual REER adjustment (i.e. the adjusted R^2s are smaller). The larger explanatory power of quarterly vs. annual misalignments is, anyhow, preserved.

Table 6. Explaining changes in actual REERs

	A. Annual model, 1980-2017									
Dependent variable: $\Delta reer_{i(t\text{-}t\text{-}5)}$	CPI-based	PPP-based	GDP deflator-based	PPI-based	ULCT-based	CPI-based	PPP-based	GDP deflator-based	PPI-based	ULCT-based
BS proxy:	GDP per capita					Labour productivity				
$Misalignment_{i(t\text{-}5)}$	-0.758***	-0.378***	-0.681***	-0.923***	-0.692***	-0.757***	-0.375***	-0.674***	-0.926***	-0.695***
	(0.020)	(0.015)	(0.020)	(0.014)	(0.031)	(0.021)	(0.015)	(0.021)	(0.013)	(0.030)
$\Delta fundam_{i(t\text{-}t\text{-}5)}$	1.984***	1.905***	1.799***	1.552***	1.782***	2.095***	1.746***	1.812***	1.897***	1.727***
	(0.116)	(0.100)	(0.097)	(0.124)	(0.215)	(0.127)	(0.098)	(0.100)	(0.162)	(0.206)
Observations	1686	1686	1686	1597	918	1629	1629	1629	1550	917
Adjusted R^2	0.537	0.373	0.502	0.755	0.394	0.521	0.365	0.490	0.774	0.396
	B. Annual model, 1999-2017									
Dependent variable: $\Delta reer_{i(t\text{-}t\text{-}5)}$	CPI-based	PPP-based	GDP deflator-based	PPI-based	ULCT-based	CPI-based	PPP-based	GDP deflator-based	PPI-based	ULCT-based
BS proxy:	GDP per capita					Labour productivity				
$Misalignment_{i(t\text{-}5)}$	-0.851***	-0.468***	-0.742***	-0.914***	-0.693***	-0.850***	-0.470***	-0.743***	-0.919***	-0.695***
	(0.028)	(0.025)	(0.029)	(0.030)	(0.035)	(0.029)	(0.025)	(0.029)	(0.030)	(0.035)
$\Delta fundam_{i(t\text{-}t\text{-}5)}$	1.430***	1.581***	1.555***	1.351***	1.451***	1.464***	1.417***	1.523***	1.644***	1.553***
	(0.152)	(0.142)	(0.138)	(0.165)	(0.265)	(0.161)	(0.136)	(0.140)	(0.218)	(0.259)
Observations	1042	1042	1042	1004	707	1040	1040	1040	1002	707
Adjusted R^2	0.476	0.303	0.435	0.484	0.362	0.476	0.295	0.432	0.489	0.367
	C. Quarterly model, 1999-2017									
Dependent variable: $\Delta reer_{i(t\text{-}t\text{-}20)}$	CPI-based	PPP-based	GDP deflator-based	PPI-based	ULCT-based					
BS proxy:	GDP per capita									
$Misalignment_{i(t\text{-}5)}$	-0.838***	-0.875***	-0.824***	-0.924***	-0.866***					
	(0.011)	(0.009)	(0.010)	(0.012)	(0.011)					
$\Delta fundam_{i(t\text{-}t\text{-}5)}$	0.962***	0.970***	0.965***	0.955***	0.890***					
	(0.021)	(0.016)	(0.017)	(0.026)	(0.035)					
Observations	3353	3353	3353	3180	2318					
Adjusted R^2	0.672	0.764	0.725	0.670	0.736					

Notes: Panel fixed effects regressions. Country fixed effects are included, but here not reported. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The estimation periods are reported in the table for each panel.

6. Conclusions

In this paper we have focused on one possible empirical approach to estimate REER misalignments, namely the BEER methodology. We have further narrowed down the analysis to assess the possible impact of employing different data frequencies in the measurement of these imbalances; to our knowledge, no study has thus far specifically tackled this relevant modelling choice.

When estimating a BEER model, the researcher is indeed faced with a potential conundrum. On the one hand, it is preferable to consider an as wide time-span as possible in order to alleviate several technical assumption underlying the BEER methodology. Yet long-term data for a satisfactory country sample is only available at an annual frequency, and with a significant lag (one year) for early-warning purposes. On the other hand, given the high-frequency nature of exchange rates it is useful to monitor these imbalances at an infra-annual frequency and in a timely fashion (i.e. one or two quarter lags relative to the reference period), thereby resorting to a quarterly BEER model, which however restricts the time-span to the last twenty years.

In our view, the findings presented in this paper are reassuring. Indeed, a state-of-the art annual BEER model has been set up and estimated over the 1980-2017 period for 55 countries: the explanatory variables have been selected according to BMA techniques, various data sources have been explored to construct the explanatory variables, the most recent panel cointegration techniques have been employed to estimate the model and manifold robustness checks have been carried out. The resulting misalignments estimates for the four main euro-area countries, taken as an insightful case-study, are found to be not systematically different from those obtained from a comparable model, based on quarterly data and estimated over the shorter, 1999Q1-2017Q4, period, and similar to that currently in use at the Bank of Italy. Some discrepancies do arise for certain countries and years, but can be explained by large movements in the (relative) investment rate and in demographic trends, which are not statistically significant in the quarterly model. Moreover, REER misalignments are found to be better predictors of subsequent REER developments in all countries when they are measured quarterly, as opposed to annually.

In conclusion, is there an "optimal" data frequency at which to estimate BEER models? Given that there is no benchmark value of REER equilibria or misalignments, validating, let alone, testing for the superiority of one model over the other is problematic, even within the common BEER approach. What can be said, however, is that, according to their use, one set of estimates can be more appropriate than the other.[32] For example, on the one hand, for research purposes, being able to examine long-run developments in these imbalances, which are affected by slow-moving variables such as demographics, and which are produced only by an annual model, is crucial (see, for instance, the vast literature on REER misalignments and economic growth, of which Habib, Mileva and Stracca, 2017; Grekou, 2018b; Giordano, 2019 are some recent examples). On the other hand, for monitoring and policy-making purposes a quarterly model – as in Giordano (2018), currently in use at the Bank of Italy –, which produces high-frequency and timely misalignment estimates is essential, as well as satisfactorily robust, as documented in this paper.

[32] Clearly, this claim is not restricted to REER misalignment modelling, but concerns many fields in economics and statistics. For example, yearly financial accounts are useful for structural analyses, such as those studying the evolution of financial systems (e.g. Bruno, De Bonis and Silvestrini, 2012); conversely, quarterly financial accounts can be employed for business-cycle analyses (e.g. Giordano, Marinucci, and Silvestrini, 2019).

Annex A. Additional details on the BEER methodology

Table A1. An overview of BEER model studies

References	Countries	Time-span	Frequency	Explanatory variables	Deflator	Estimation methodology
Cubeddu et al. (2019)	40 (advanced and emerging economies) + EA (as aggregation)	1990-2016	A	Net foreign assets (-/+), expected GDP growth (+), public health expenditure (+), VIX and interactions (-/0), share of currency held as FX reserve by central banks (0/-), output gap (+), terms of trade (+), openness (-), private credit (+/0), change in reserves (-), population growth, old age dependency ratio (-/+), real interest rate (+), GDP (+), domestic debt owned by residents (+), share of administered prices (-), political risk rating (+), capital stock per person employed (+), VAT revenue	CPI; PPP	Fixed effects OLS
Giordano (2018)	57 (advanced and emerging economies) + EA	1999-2017	Q	GDP per capita (+); trade openness (-); terms of trade(+); government expenditure(+); real interest rate(+)	CPI; GDP deflator; PPP; PPI (Banca d'Italia); ULCT	CCEMG
Fidora, Giordano and Schmitz (2018)	57 (advanced and emerging economies) + EA	1999-2016	Q	GDP per capita/labour productivity(+); trade openness(-); terms of trade(+); government expenditure(+); real interest rate(+)	CPI; GDP deflator; PPP; PPI (ECB); ULCT	CCEMG
Couharde et al. (2017)	182 (advanced and emerging economies)	1973-2016	A	GDP per capita (+), net foreign assets (+); terms of trade (+)	CPI	PMG
Adler and Grisse (2017)	20 advanced economies	1980-2013	A	GDP per capita (+); government expenditure(+); net foreign assets (+); terms of trade(+); central bank reserves (-); old age dependency rate (+); openness (-); real interest rate (+); trade balance (-)	PPP; CPI	DOLS
Comunale (2017; 2018)	28 (EU countries)	1994-2012	A	Net foreign assets (+)/cumulated trade balances (+); terms of trade(+); GDP per capita(+)	CPI	GM-FMOLS
Gnimassoun and Mignon (2015)	22 (industrialized countries)	1980-2011	A	GDP per capita (+): net foreign assets (+)	CPI	DOLS
Ajevskis et al. (2014)	11 euro-area countries	2001-2010	Q	Labour productivity (+); trade openness (-); government consumption; investment ratio; net foreign assets (+); terms of trade (+); fiscal balance to GDP (-)	CPI	VECM
Fischer and Hossfeld (2014)	57 (advanced and emerging economies)	1980-2011	A	Labour productivity(+)	PPP	a) Fixed effects OLS b) Pooled OLS c) DOLS
Mancini-Griffolo, Meyer, Natal and Zanetti (2015)	18 (advanced economies)	1973-2011	A	Net foreign assets(+);GDP per capita(+); terms of trade(+); government consumption(+); sectorial labour productivity	CPI; PPI	DOLS
Du, Wei and Xie (2013)	61 (advanced and emerging economies)	1988-2007	A	GDP per capita(+); road density(-); government expenditure(+); real interest rate(+); labour productivity(+)	CPI	a) Fixed effects OLS; b) 2SLS
Coudert, Couharde and Mignon (2013)	11 (EA countries)	1980-2010	A	GDP per capita (+): net foreign assets (+)	CPI	DOLS

Notes: A=annual; Q=quarterly. The explanatory variables reported are those included in the baseline specifications of the selected studies. When the + or - sign is omitted the estimated relationship is not statistically significant. This table is an update and expansion of a similar table first published in Fidora, Giordano and Schmitz (2018).

Table A1 cont.

References	Countries	Time-span	Frequency	Explanatory variables	Deflator	Estimation methodology
Bussière et al. (2010)	a) 44 b) 14 (advanced and emerging economies)	1980-2007	a) A b) Q	Commodity terms of trade(+); fiscal policy(+); civil liberties(-); openness(-); net foreign assets; investment; government expenditure; trade restriction index; GDP per capita (+); commodity prices	PPP	Panel estimations: a) ARDL; b) CCEMG; c) CCEP
Hossfeld (2010)	17 (US and its 16 major trading partners)	1986-2006	Q	Net foreign assets (-); trade balance; terms of trade(+); government consumption; openness	CPI	Panel estimations: a) GM DOLS; b) FMOLS
Bénassy-Queré, Béreau and Mignon (2009; 2010)	15 (advanced and emerging economies)	1980-2005	A	CPI to PPI ratio/industry to services value added deflator/GDP per capita/labour productivity (+); net foreign assets (+); real interest rate(+); terms of trade (+)	CPI	DOLS
Bénassy-Queré, Lahrèche-Révil and Mignon (2008)	16 (advanced and emerging economies)	1980-2004		CPI to PPI ratio(+); net foreign assets(+)	CPI	DOLS
Ricci, Milesi-Ferretti and Lee (2008)	48 (advanced and emerging economies)	1980-2004	A	Trade restriction index(+); price controls(-); commodity terms of trade(+); net foreign assets to trade(+);government expenditure (+); labour productivity tradables(+); labour productivity nontradables(-)	CPI	a) DOLS b) FMOLS
Aguirre and Caldéron (2005)	60 (advanced and emerging economies)	1963-2003	A	GDP per capita (+); terms of trade (+); government spending (+); net foreign assets (+)	CPI	a) DOLS b) PDOLS
Lane and Milesi-Ferretti (2004)	64 (advanced and emerging economies)	1975-1996	A	Net foreign assets (+); GDP per capita (+); terms of trade (+)	CPI; WPI	DOLS
Maeso Fernández, Osbat and Schnatz (2004)	25 (OECD countries)	1975-2002	A	GDP per capita(+); government expenditure to GDP(+); openness(-)	PPP	a) MGE /PMGE b) FMOLS c) DOLS
Nilsson (2004)	15 (OECD countries)	1982-2000	Q	Terms of trade (+); CPI to PPI ratio (+); net foreign assets (+); relative stock of government debt (+); real interest rate (-)	CPI	VAR(2)
Maeso Fernández, Osbat and Schnatz (2001)	23 (advanced economies)	1975-1998	Q	Labour productivity (+); accumulated current account to GDP; real price of oil (+); real interest rate (-)	CPI	VECM
Clark and MacDonald (1998)	7 (G-7 countries)	1960-1996	A	Terms of trade (+); CPI to PPI ratio (+); net foreign assets (+); relative stock of government debt (+); real interest rate (-)	CPI	VECM

Notes: A=annual; Q=quarterly. The explanatory variables reported are those included in the baseline specifications of the selected studies. When the + or - sign is omitted the estimated relationship is not statistically significant. This table is an update and expansion of a similar table first published in Fidora, Giordano and Schmitz (2018).

The choice of the dependent variable

The BEER models described in this paper use the bilateral RER, as opposed to the REER, as the dependent variable for the reasons discussed in Section 3. Moreover, bilateral exchange rates turn out to be significantly cointegrated with bilaterally specified economic fundamentals.

However, one could argue that the theoretical exchange rate that is consistent with both internal and external equilibrium is the REER, not the RER (e.g. Chinn, 2006). Our approach is therefore exactly the opposite to that of deriving bilateral equilibria and misalignments from their effective counterparts, as in, for example, Alberola et al. (1999) and Bénassy-Quéré, Béreau and Mignon (2008; 2009). The indirect derivation of bilateral RERs developed in these studies is based on the inversion of the weighted matrix of effective equilibrium exchange rates. Since only (N-1) independent bilateral exchange rates can be derived from N effective rates (the so-called "redundancy problem"; Faruqee, 1998), it, however, implies dropping one of the currencies, corresponding to the numeraire currency, and thus rests on the assumption that the misalignment of the numeraire currency is the mirror image of all other (N-1) countries' misalignments.

To alleviate this shortcoming, Faruqee, Isard and Masson (1999) and Bénassy-Quéré, Lahrèche-Révil and Mignon (2011) use the "rest of the world" (RoW) aggregate as the numeraire: REERs are calculated for a sample of (N+1) countries, including the RoW, and N bilateral misalignments are then derived against the RoW. This approach is, however, appropriate only to the extent that the RoW misalignment (in effective terms) is the mirror image of that of the N countries of the sample. Coudert et al. (2019) adopt a different approach. In particular, they decompose the equilibrium REER into the weighted product of all countries' bilateral equilibrium nominal exchange rates (NERs) multiplied by their relative prices (similarly to the decomposition of actual REERs). By using the no-arbitrage condition on the foreign exchange market that guarantees the consistency of all cross rates, they next express the bilateral equilibrium NER of each country relative to a numeraire currency, namely Special Drawing Rights (SDR). Frankel and Wei (2008) too argue in favour of using SDRs since they represent a weighted index of currencies. The use of SDRs thus allows defining the value of each currency independently from the other currencies and, in turn, to derive a bilateral equilibrium exchange rate path *vis-à-vis* the SDR for each country that is consistent with minimized currency misalignments (i.e. when the actual REER is equal to the equilibrium REER).

Another alternative is to correct for multilateral consistency, as in Adler and Grisse (2017) and Cubeddu et al. (2019), following Faruqee (1998). The latter study suggests a procedure for rectifying the problem by first observing that one of the eigenvalues of the time-specific matrix of trade weights underlying the REER must equal unity because the columns of this matrix sum to unity. Because "first-stage" REER misalignments do not guarantee that the weighted average of residuals of all countries is zero in each year, REERs are adjusted by the global weighted average of residuals. For each year the weights are given by the eigenvector associated with the unit eigenvalue of the trade weights matrix for that year.

By estimating bilateral equilibria directly instead, as in this paper, these assumptions or adjustments can be avoided. The numeraire country's bilateral misalignment *vis-à-vis* a competitor is the mirror of the competitor's bilateral misalignment *vis-à-vis* the numeraire, which can then be aggregated up to an effective misalignment by using the numeraire's appropriate trade weights. Our approach, however, naturally does not avoid having to select a numeraire country, a choice which is discussed in Section 3.

The measurement of trade costs

As a proxy of trade costs, in this paper we mainly employ the so-called "phi-ness" indicator of trade, which derives the level of trade impediments from bilateral trade flows. Starting from the structural gravity specification in Anderson and van Wincoop (2004), nominal bilateral exports X_{ij} from exporter country i to importer country j, where time subscripts are suppressed for simplicity, can be expressed as:

(A1) $X_{ij} = \frac{Y_i E_j}{Y} \left(\frac{t_{ij}}{\Pi_i P_j}\right)^{(1-\sigma)}$,

in turn the product of two terms. As explained in Yotov et al. (2016), the first is a size term, which includes the output Y_i of the exporter i and the expenditure E_j of the partner j relative to global output Y, and represents the hypothetical level of frictionless trade between partners i and j if there were no trade costs. Under this scenario, consumers face the same price for a given variety (i.e. goods from different countries) regardless of their physical location and their expenditure share on goods from a given country is equal to the share of production in the exporter country in the global economy (i.e. $\frac{X_{ij}}{E_j} = \frac{Y_i}{Y}$). This term suggests that large producers will export more to all destinations, large or rich markets will import more from all sources and trade flows between partners *i* and *j* will be larger the more similar in size the two countries are. The second term is the trade cost term, which in turn consists of three components: bilateral trade costs between partners *i* and *j* (t_{ij}), the structural term P_j, which measures inward multilateral resistance, that is importer *j*'s ease of market access, and the structural term Π_i, i.e. the outward multilateral resistance, which measures exporter i's ease of market access. Since the elasticity of substitution among different varieties σ is larger than one, a rise in bilateral trade costs is found to reduce bilateral trade flows; conversely, larger multilateral resistance terms lead to higher trade.

As discussed in Chen and Novy (2012), although most researchers opt for a log-linear trade cost function, the estimated effect on trade costs depends on the chosen functional form. Moreover, as many trade costs are unobservable, a potential omitted variable bias is a concern. Finally, many trade cost proxies have no time variation, which makes it difficult to track changes in trade costs. These issues may be overcome by indirectly inferring implied trade costs from bilateral trade data without specifying a trade cost function. The idea is to isolate t_{ij} from the gravity equation (A1) and express it in terms of observable trade data. However, the multilateral resistance variables are theoretical constructs and are therefore unobservable. As shown in Head and Ries (2001) and Novy (2013), these terms can however be eliminated by multiplying equation (A1) by its counterpart for trade flows in the opposite direction, X_{ji}, and then dividing it by the product of gravity equations for domestic trade flows within each country $X_{ii}X_{jj}$. Taking the square root results in the expression for the phi-ness of trade, ϕ_{ij}:

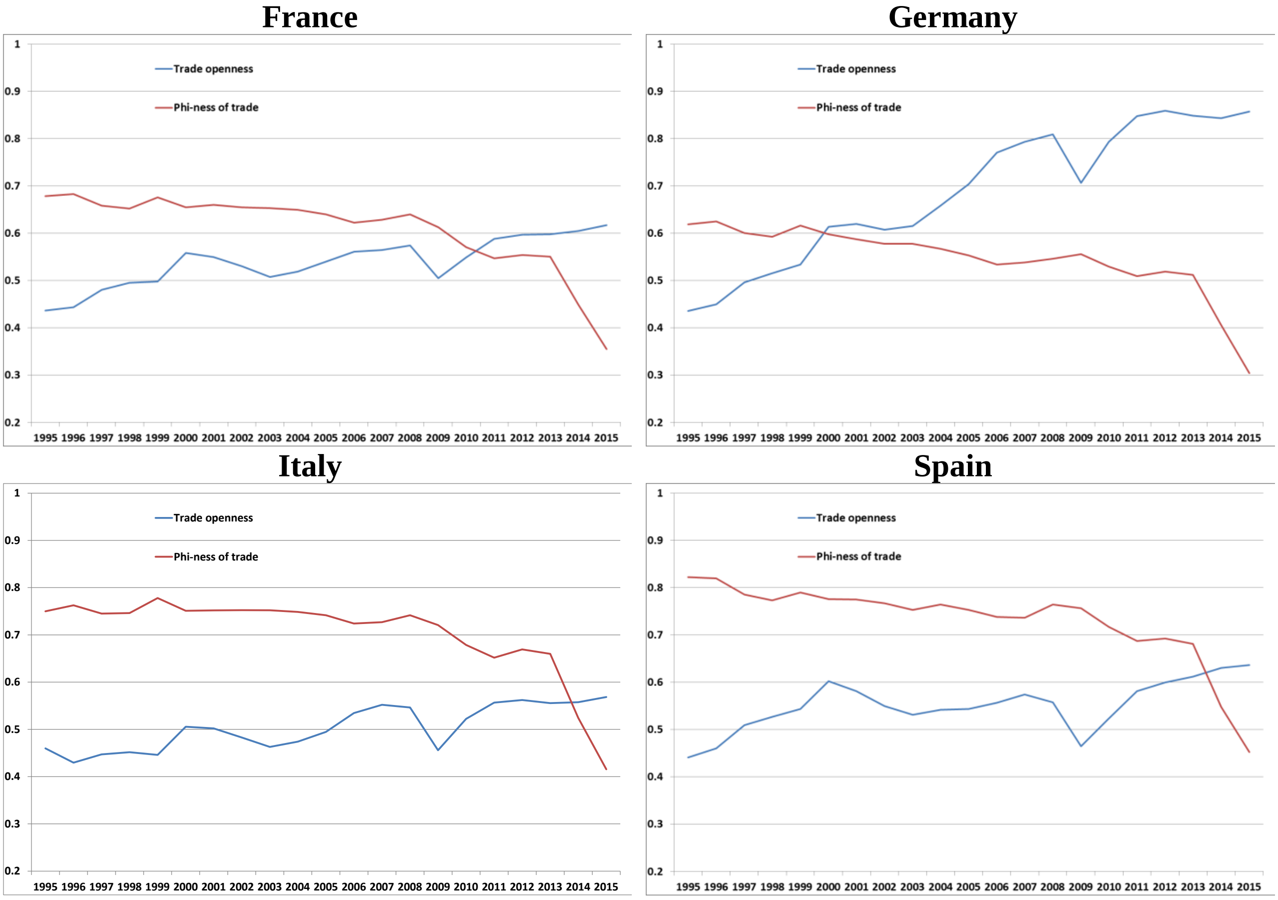
(A2) $\phi_{ij} = (\frac{X_{ij}X_{ji}}{X_{ii}X_{jj}})^{\frac{1}{2}} = (\frac{\tau_{ij}\tau_{ji}}{\tau_{ii}\tau_{jj}})^{\frac{(1-\sigma)}{2}}$

This indicator measures bilateral trade costs relative to domestic trade costs and is by construction symmetric, as it is the geometric average of trade costs in both directions. It can be constructed using bilateral trade data. However, domestic trade data, X_{ii} and X_{jj}, are hard to find, and are usually proxied by apparent consumption, i.e. by deducting exports from total domestic production (see, for example, Wei, 1996 and Novy, 2013 for a discussion).

The measure in equation (A2) employed in this paper is taken from the ESCAP-WB Trade Cost Database, described in Arvis et al. (2012), for the years for which it is available (1995-2015). As it is a bilateral measure, we construct an effective measure by weighting it with trade weights sourced from the ECB.

Figure A1 plots both the (effective) phi-ness indicator and openness to trade of the four main euro-area countries: the correlation between the two series is high over the 1995-2015 period and ranges between -0.65 (Italy) and -0.75 (Germany).

Figure A1. Phi-ness of trade and trade openness in the four main euro-area countries
(percentages)



Sources: author’s calculations on ESCAP-WB trade, ECB and IMF-WEO data.

The selection of explanatory variables via Bayesian model averaging (BMA)

In BMA the parameters of a given model are treated as random, distributed according to a prior distribution. The variable indicating whether a given model is true is also treated as random and distributed according to a prior distribution.

As in Magnus, Powell, and Prüfer (2010), the general set-up of BMA is the standard linear regression model, such that the models differ solely by which subset of predictors are contained in each model. In particular,:

(A3) $y = X_1\beta_1 + X_2\beta_2 + \varepsilon,$

where $y(nx1)$ is the vector of observations, $X_1(nxk_1)$ and $X_2(nxk_2)$ are matrices of regressors, and ε is a random vector of unobservable disturbances. We assume that $k_1 \geq 1$, $k_2 \geq 0$, $k = k_1 + k_2 \leq (n-1)$ and that the disturbances are i.i.d $\sim N(0, \sigma^2)$. X_1 contains the “focus” regressors, that necessarily need to be included in the model, for example on theoretical grounds, irrespective of their statistical significance. X_2 contains the “auxiliary” regressors, whose inclusion needs to be tested. There are k_2 components of β_2 and a different model arises whenever a different subset of the parameters β_2 are set equal to zero. In other terms, there are 2^{k_2} possible combinations of auxiliary regressors, i.e. 2^{k_2} different models which all seek to explain D, the data.

The jth model M_j is the following:

(A4) $y = X_1\beta_1 + X_{2j}\beta_{2j} + \varepsilon$

where X_{2j} is the (nxk_{2j}) matrix containing a subset of k_{2j} columns of X_2, β_{2j} is the corresponding $(k_{2j}x1)$ subvector of β_2 and $0 \leq k_{2i} \leq k_2$. Model averaging estimation proceeds in two steps. The first step requires how to estimate the parameters, conditional upon a selected model. In the second step the estimator is computed as a weighted average of these conditional estimators. Under BMA, Bayesian techniques are employed in both stages.

Following Moral-Benito (2012) and the Bayesian logic, the posterior for the parameters calculated using M_j is written as:

(A5) $(\beta_{2j}|D, M_j) = \frac{f(D|\beta_{2j},M_j)\ g(\beta_{2j}\ |M_j)}{f(D\ |M_j)}$

such that for each model there is a posterior $g(\beta_{2j}\ |M_j)$, a likelihood $f(D|\beta_{2j}, M_j)$ and a prior $f(D\ |M_j)$. Given a prior model probability $P(M_j)$, the posterior model probability, which summarises model uncertainty, can be calculated using Bayes' Rule as:

(A6) $P(M_j|D) = \frac{f(D|\ M_j)\ P(M_j)}{f(D\ )}$.

$P(M_j)$, which does not depend on D, measures how likely we believe M_j to be the correct model before turning to the data. The integrated likelihood $f(D|\ M_j)$, also called marginal probability of the data since it is obtained by integrating the joint density of (D, β_{2j}) given D over β_{2j}, can be derived by integrating both sides of (A5) with respect to β_{2j}, exploiting the fact that probability density functions integrate to 1 (i.e. $\int g(\beta_{2j}|D, M_j)\ d\beta_{2j} = 1$), and rearranging the terms such that:

(A7) $f(D|\ M_j) = \int f(D|\beta_{2j}, M_j)\ g(\beta_{2j}\ |M_j)\ d\beta_{2j}$.

The ratio of the integrated likelihoods of two different models, the so-called Bayes Factor, is similar to the likelihood ratio statistics, in which the parameters β_{2j} are eliminated by maximization instead of by integration.

Furthermore, considering β_2 a function of β_{2j}, for each j=1…2^{k_2}, and by the law of total probability, the posterior density of the parameters for all models under consideration is:

(A8) $g(\beta_2|D) = \sum_{j=1}^{2^{k_2}} P(M_j|D)\ \ g(\beta_2\ |D, M_j)$,

that is the weighted sum of the posterior density of each model, with weights in the averaging process being their posterior model probability.

The point estimators of the parameters can be obtained by taking expectations across (A8). As in Leamer (1978), the posterior variance can instead be computed as:

(A9) $V(\beta_2|D) = \sum_{j=1}^{2^{k_2}} P(M_j|D)\ \ V(\beta_2\ |D, M_j) + \sum_{j=1}^{2^{k_2}} P(M_j|D)\ \ (E(\beta_2\ |D, M_j) - E(\beta_2\ |D))^2$

such that it includes both the weighted average of the estimated variances of the individual models and the weighted variance in estimates of the β_2's across different models. This implies that even if estimates are very precise in all the models, there may be considerable uncertainty surrounding the parameter if these estimates are very different across specifications.

Finally, one can estimate the posterior inclusion probability (PIP), i.e. the posterior probability that a particular auxiliary variable h is included in the regression, and therefore belongs to the true model. It is calculated as the sum of the posterior model probabilities for all of the models including that variable:

(A10) PIP $= P(\beta_{2h} \neq 0|D) = \sum_{\beta_{2h}\neq 0} P(M_j|D)$.

As mentioned earlier, implementation of BMA requires choosing two sets of priors, the prior distribution of the parameters given the model and the prior probability of the model. In Magnus, Powell, and Prüfer (2010)'s setup, given the normal regression framework, and following Fernandez, Ley and Steel (2001), prior beliefs on the regression parameters of model M_i are introduced according to a hierarchical prior structure by imposing: a) conventional non-informative priors on the common parameters β_1 and the error variance σ^2;[33] and b) an informative fixed g-prior specification of Zellner (1986) on the auxiliary parameters β_{2i}, specifically $\beta_{2i}|\beta_1, \sigma^2, M_i \sim$ N(0, $\sigma^2 V_{0i}$), where ${V_{0i}}^{-1} = g\boldsymbol{X}_{2i}^T M_i \boldsymbol{X}_{2i}$ and $g = \frac{1}{\max(n, k_2^2)}$ is a constant scalar for each model M_i.[34] In other terms, for the auxiliary parameters a normal density with zero mean and a prior variance that is proportional to the posterior covariance of the sample is assumed. The g scalar thus defined is also known as the "Benchmark Prior" and it determines how much importance is attributed to the prior beliefs of the researcher. Indeed, it captures the uncertainty related to the coefficients being zero: a small g implies a small coefficient variance and a higher confidence on the coefficient being zero; the opposite is true when g is large, such that higher values of g correspond to greater shrinkage. Finally, equal prior probability is assigned to each model, so that $p(M_i) = 2^{-k_2}$ (i.e. uniform prior on the model space). This prior structure is a special case of the binomial prior, in which the probability of success parameter is equal to ½. It implicitly assumes that the probability of one regressor appearing in the model is independent of the inclusion of others; for this reason alternative proxies of the same variable (such as the BS effect, or demographics) are not included jointly, as discussed in footnote 11.

In Table A2 we report the PIPs for each auxiliary variable in an alternately-deflated BEER model.

[33] In other terms, given the absence of information on these unknown coefficients, we assume complete uncertainty where the prior is located (i.e. improper priors that do not integrate to one).

[34] As discussed in Moral-Benito (2015), to which we refer for a comprehensive overview on different prior structures used in the context of model uncertainty, the popularity of this prior structure is due to two factors: (i) it has closed-form solutions for the posterior distributions, which drastically reduce the computational burden, and (ii) it only requires the elicitation of one hyperparameter, the scalar g. Various approaches to choosing g have been proposed in the literature. The Unit Information Prior (g-UIP), proposed by Kass and Wasserman (1995), corresponds to taking g=n, and it leads to Bayes factors that behave like the BIC. The Risk Inflation Criterion (g-RIC) implies setting $g = k_2^2$. The Benchmark Prior, used in this paper, is a combination of the g-UIP and g-RIC priors, which has been found to perform best as concerns predictive performance.

Table A2. Selecting the annual economic fundamentals via BMA

GDP per capita as the proxy of the BS effect

CPI-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-1.194	0.087	-13.710	1.00	-1.281	-1.107
Relative GDP per capita	1.186	0.099	12.010	1.00	1.087	1.285
Relative trade cost	-0.521	0.042	-12.380	1.00	-0.563	-0.479
Relative terms of trade	-0.014	0.090	-0.150	0.04	-0.104	0.077
Relative government consumption	12.199	1.220	10.000	1.00	10.979	13.418
Relative real interest rate	0.000	0.000	0.170	0.05	0.000	0.000
Relative investment rate	-5.341	0.957	-5.580	1.00	-6.298	-4.383
Relative net foreign assets	0.000	0.001	-0.070	0.03	-0.001	0.001
Relative old-age dependency ratio	-4.461	1.071	-4.170	1.00	-5.531	-3.390
Relative credit to the private sector	-0.025	0.287	-0.090	0.03	-0.311	0.262
Number of observations	*1335*					

PPI-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-1.128	0.090	-12.520	1.00	-1.218	-1.038
Relative GDP per capita	1.223	0.102	12.010	1.00	1.121	1.325
Relative trade cost	-0.525	0.043	-12.300	1.00	-0.567	-0.482
Relative terms of trade	-0.005	0.061	-0.080	0.03	-0.065	0.056
Relative government consumption	11.753	1.229	9.560	1.00	10.524	12.982
Relative real interest rate	0.000	0.000	0.150	0.04	0.000	0.000
Relative investment rate	-5.907	0.972	-6.080	1.00	-6.879	-4.935
Relative net foreign assets	0.000	0.001	-0.040	0.03	-0.001	0.001
Relative old-age dependency ratio	-4.934	1.067	-4.620	1.00	-6.001	-3.866
Relative credit to the private sector	-0.028	0.302	-0.090	0.03	-0.330	0.274
Number of observations	*1320*					

PPP-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-0.587	0.017	-34.110	1.00	-0.605	-0.570
Relative GDP per capita	0.262	0.018	14.230	1.00	0.244	0.281
Relative trade cost	0.000	0.001	-0.050	0.03	-0.001	0.001
Relative terms of trade	0.710	0.055	12.910	1.00	0.655	0.765
Relative government consumption	2.223	0.218	10.220	1.00	2.006	2.441
Relative real interest rate	0.000	0.000	0.180	0.05	0.000	0.000
Relative investment rate	0.002	0.035	0.070	0.03	-0.033	0.038
Relative net foreign assets	0.000	0.000	0.050	0.03	0.000	0.000
Relative old-age dependency ratio	1.718	0.188	9.120	1.00	1.530	1.906
Relative credit to the private sector	4.473	0.240	18.610	1.00	4.233	4.713
Number of observations	*1335*					

ULCT-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-1.440	0.105	-13.720	1.00	-1.545	-1.335
Relative GDP per capita	0.287	0.191	1.500	0.77	0.096	0.478
Relative trade cost	-0.701	0.042	-16.610	1.00	-0.744	-0.659
Relative terms of trade	3.928	0.423	9.280	1.00	3.505	4.351
Relative government consumption	2.814	1.933	1.460	0.75	0.881	4.747
Relative real interest rate	-0.064	0.024	-2.680	0.95	-0.088	-0.040
Relative investment rate	-7.568	1.356	-5.580	1.00	-8.924	-6.213
Relative net foreign assets	-0.010	0.010	-0.960	0.55	-0.020	0.000
Relative old-age dependency ratio	0.093	0.509	0.180	0.07	-0.416	0.602
Relative credit to the private sector	0.008	0.309	0.030	0.04	-0.301	0.317
Number of observations	*778*					

GDP-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-1.197	0.087	-13.820	1.00	-1.283	-1.110
Relative GDP per capita	1.179	0.098	12.020	1.00	1.081	1.277
Relative trade cost	-0.503	0.042	-12.000	1.00	-0.544	-0.461
Relative terms of trade	0.006	0.064	0.100	0.03	-0.058	0.070
Relative government consumption	12.631	1.212	10.420	1.00	11.418	13.843
Relative real interest rate	0.000	0.000	0.210	0.06	0.000	0.000
Relative investment rate	-5.266	0.953	-5.520	1.00	-6.219	-4.313
Relative net foreign assets	0.000	0.001	-0.060	0.03	-0.001	0.001
Relative old-age dependency ratio	-4.574	1.055	-4.340	1.00	-5.629	-3.519
Relative credit to the private sector	-0.012	0.249	-0.050	0.03	-0.261	0.237
Number of observations	*1335*					

Notes: Each panel refers to an alternately deflated RER. Estimation is conducted via the BMA estimator introduced by Magnus, Powell, and Prüfer (2010) on the country sample excluding the US. The model space is equal to 512 models in all cases. Country fixed effects are the only focus regressors, whereas all other variables are considered as auxiliary covariates. The variables with low (i.e. under 0.5) posterior probability of inclusion are highlighted in red. t-stats correspond to the posterior mean divided by the posterior standard deviation.

Table A2. continued

Labour productivity as the proxy of the BS effect

CPI-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-1.302	0.082	-15.910	1.00	-1.384	-1.220
Relative labour productivity	1.184	0.103	11.490	1.00	1.081	1.288
Relative trade cost	-0.503	0.042	-12.030	1.00	-0.545	-0.461
Relative terms of trade	0.012	0.083	0.140	0.04	-0.072	0.095
Relative government consumption	10.591	1.173	9.030	1.00	9.418	11.764
Relative real interest rate	0.000	0.000	0.260	0.09	0.000	0.000
Relative investment rate	-4.094	1.025	-3.990	1.00	-5.119	-3.069
Relative net foreign assets	0.000	0.001	-0.070	0.03	-0.001	0.001
Relative old-age dependency ratio	-0.286	0.768	-0.370	0.15	-1.054	0.482
Relative credit to the private sector	0.156	0.663	0.230	0.08	-0.507	0.818
Number of observations	*1322*					

PPI-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-1.278	0.083	-15.320	1.00	-1.361	-1.194
Relative labour productivity	1.149	0.107	10.780	1.00	1.043	1.256
Relative trade cost	-0.502	0.043	-11.750	1.00	-0.545	-0.460
Relative terms of trade	0.029	0.134	0.220	0.07	-0.105	0.162
Relative government consumption	10.178	1.221	8.340	1.00	8.958	11.399
Relative real interest rate	0.000	0.000	0.240	0.08	0.000	0.000
Relative investment rate	-4.552	1.011	-4.500	1.00	-5.564	-3.541
Relative net foreign assets	0.000	0.001	-0.040	0.03	-0.001	0.001
Relative old-age dependency ratio	-0.392	0.904	-0.430	0.19	-1.295	0.512
Relative credit to the private sector	0.144	0.642	0.220	0.07	-0.498	0.786
Number of observations	*1307*					

PPP-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-0.653	0.015	-42.970	1.00	-0.668	-0.638
Relative labour productivity	0.235	0.019	12.180	1.00	0.216	0.254
Relative trade cost	0.000	0.002	-0.080	0.03	-0.002	0.001
Relative terms of trade	0.816	0.056	14.650	1.00	0.760	0.871
Relative government consumption	2.146	0.224	9.590	1.00	1.922	2.370
Relative real interest rate	0.000	0.000	-0.080	0.03	0.000	0.000
Relative investment rate	0.021	0.092	0.230	0.07	-0.071	0.114
Relative net foreign assets	0.000	0.000	0.050	0.03	0.000	0.000
Relative old-age dependency ratio	2.252	0.177	12.730	1.00	2.075	2.429
Relative credit to the private sector	5.325	0.239	22.320	1.00	5.086	5.564
Number of observations	*1322*					

ULCT-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-1.499	0.099	-15.070	1.00	-1.598	-1.400
Relative labour productivity	0.057	0.128	0.450	0.21	-0.071	0.186
Relative trade cost	-0.712	0.043	-16.710	1.00	-0.754	3 -.6692694
Relative terms of trade	4.108	0.411	10.010	1.00	3.697	4.518
Relative government consumption	2.286	1.958	1.170	0.65	0.328	4.244
Relative real interest rate	-0.060	0.026	-2.290	0.91	-0.086	-0.034
Relative investment rate	-8.294	1.326	-6.250	1.00	-9.621	-6.968
Relative net foreign assets	-0.009	0.010	-0.900	0.51	-0.019	0.001
Relative old-age dependency ratio	0.155	0.634	0.240	0.09	-0.479	0.789
Relative credit to the private sector	0.038	0.360	0.110	0.04	-0.321	0.398
Number of observations	*778*					

GDP-deflated RER	Coef.	Std. Err.	t	Posterior inclusion probability	[1-Std. Err. Bands]	
Constant	-1.332	0.084	-15.870	1.00	-1.415	-1.248
Relative labour productivity	1.143	0.103	11.120	1.00	1.040	1.246
Relative trade cost	-0.492	0.042	-11.640	1.00	-0.535	5 -.4501718
Relative terms of trade	0.337	0.433	0.780	0.43	-0.095	0.770
Relative government consumption	10.859	1.187	9.150	1.00	9.671	12.046
Relative real interest rate	0.000	0.000	0.190	0.06	0.000	0.000
Relative investment rate	-4.016	1.054	-3.810	0.99	-5.070	-2.962
Relative net foreign assets	0.000	0.001	-0.060	0.03	-0.001	0.001
Relative old-age dependency ratio	-0.267	0.747	-0.360	0.15	-1.015	0.480
Relative credit to the private sector	0.259	0.868	0.300	0.11	-0.609	1.127
Number of observations	*1322*					

Notes: Each panel refers to an alternately deflated RER. Estimation is conducted via the BMA estimator introduced by Magnus, Powell, and Prüfer (2010) on the country sample excluding the US. The model space is equal to 512 models in all cases. Country fixed effects are the only focus regressors, whereas all other variables are considered as auxiliary covariates. The variables with low (i.e. under 0.5) posterior probability of inclusion are highlighted in red. t-stats correspond to the posterior mean divided by the posterior standard deviation.

The estimation procedure

A1. Pesaran's (2015) weak cross-sectional dependence (CSD) test [35]

As in Chudik and Pesaran (2015), consider the following panel data model for $i = 1, \ldots, N$ and $t = 1, \ldots, T$:

(A11) $y_{it} = \alpha_i + \boldsymbol{\beta}_i' \boldsymbol{x}_{it} + u_{it}$

where $\boldsymbol{x}_{it}$ is a (k+1)x1 vector of observed regressors and the fixed unknown coefficients $\boldsymbol{\beta}_i$ are allowed to vary across i. For each i, $u_{it} \sim iid(0, \sigma_i{}^2)$ for all t, although they could be cross-sectionally correlated. Let $\hat{u}_{it}$ be the OLS estimator of u_{it} defined as:

(A12) $\hat{u}_{it} = y_{it} - \hat{\alpha}\hat{\boldsymbol{\beta}}_i' \boldsymbol{x}_{it}$

[35] See Chudik and Pesaran (2015, pp. 6-7) for a formal definition of weak vs. strong CSD.

with $\hat{a}_i$ and $\hat{\boldsymbol{\beta}}_i$being the OLS estimates of α_i and $\boldsymbol{\beta_i}$. Define $\hat{\rho}_{ij}$ as the sample estimate of the pairwise correlation of the residuals $\hat{u}_{it}$ and $\hat{u}_{jt}$, specifically:

(A13) $\hat{\rho}_{ij}=\hat{\rho}_{ji} = \frac{\sum_{t=1}^{T}\hat{u}_{it}\hat{u}_{jt}}{(\sum_{t=1}^{T}\hat{u}_{it}^2)^{1/2}(\sum_{t=1}^{T}\hat{u}_{jt}^2)^{1/2}}$. Pesaran (2015) bases the test of weak CSD on its sample estimate, given by:

(A14) $\hat{\rho}_N = \frac{2}{N(N-1)}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}$.

The CSD statistic can be written as:

(A15) $CSD = [\frac{TN(N-1)}{2}]^{1/2}\hat{\rho}_N$.

Under the null hypothesis, CSD$\sim N(0,1)$.

Table A3 provides the results of the weak CSD test for all the variables included in the annual BEER model.

Table A3. Testing for CSD in the annual panel dataset

	Pesaran's (2015) test of weak CSD	P-value
GDP per capita	14.829	0.000
Labour productivity	17.757	0.000
Trade cost	38.461	0.000
Terms of trade	45.805	0.000
Government consumption	87.706	0.000
Investment rate	23.519	0.000
Old-age dependency ratio	27.818	0.000
CPI-based RER	63.653	0.000
GDP deflator-based RER	72.944	0.000
PPP-based RER	117.724	0.000
PPI-based RER	57.209	0.000
ULCT-based RER	27.677	0.000

Notes: H0: weak CSD; H1: strong CSD. The US is excluded from the sample.

A2. Panel unit root tests

In order to correct standard unit root tests for the presence of CSD, Pesaran (2007) proposes augmenting standard augmented Dickey–Fuller (ADF) regressions with the cross-section averages of lagged levels and first-differences of the individual series. Standard panel unit root tests are then based on the simple averages of the individual cross-sectionally augmented ADF statistics.

In more detail, suppose that y_{it} for $i = 1, \ldots, N$ and $t = 1, \ldots, T$ is generated according to a simple dynamic linear heterogeneous panel data model:

(A16) $y_{it} = (1-\varphi_i)\mu_i + \varphi_i y_{it-1} + u_{it}$

where the error term u_{it} has a single-factor structure, implying a single unobserved common factor:

(A17) $u_{it} = \gamma_i f_t + \varepsilon_{it}$

in which f_t is the common unobserved factor, γ_i is the country-specific factor loading and ε_{it} is the individual-specific error. Equations (A15) and (A16) can be rewritten as follows:

(A18) $\Delta y_{it} = \alpha_i + \beta_i y_{it-1} + \gamma_i f_t + \varepsilon_{it}$

where $\alpha_i = (1-\varphi_i)\mu_i$, $\beta_i = -(1-\varphi_i)$ and $\Delta y_{it} = y_{it} - y_{it-1}$. The unit root hypothesis, $\varphi_i = 1$, can thus be expressed as: H_0: $\beta_i = 0$ for all *i* against the alternative hypothesis of H_1: $\beta_i < 0 \; for \; i = 1, \dots, N_1,\ \beta_i = 0 \; for \; i = N_1 + 1, \dots,$ N. Let $\bar{\gamma} = N^{-1}\sum_{j=1}^{N}\gamma_j$. Then the common factor f_t can be proxied by the cross-section mean of y_{it}, namely $\bar{y_t} = N^{-1}\sum_{j=1}^{N} y_{jt}$ and its lagged values $\overline{y_{t-1}}, \overline{y_{t-2}}$, … In the simple case when u_{it} is serially uncorrelated, $\bar{y_t}$ and $\overline{y_{t-1}}$ are sufficient for asymptotically filtering out the effects of the common unobserved factor f_t. Thus, the test of the unit root hypothesis can be based on the t-ratio of the OLS estimate $\widehat{b_i}$ in the following cross-sectionally ADF (CADF) regression:

(A19) $\Delta y_{it} = a_i + b_i y_{it-1} + c_i \overline{y_{t-1}} + d_i \Delta y_t + e_{it}$.

The cross-sectionally augmented version of the IPS (Im, Pesaran and Shin, 2003) test becomes:

(A20) $CIPS(N,T) = N^{-1}\sum_{j=1}^{N} t_i(N,T)$

where $t_i(N,T)$ is the CADF statistic for the *i*th cross-section unit given by the t-ratio of the coefficient of y_{it-1} in the CADF regression defined by (A18). The asymptotic distribution of CIPS converges to N(0,1).

Table A4 provides the CIPS panel unit root test results for all the variables included in the annual BEER model, with and without a time trend.

Table A4. Unit-root testing in the annual panel dataset

(p-values)

	Pesaran's (2007) CIPS test - with trend	Pesaran's (2007) CIPS test - without trend
GDP per capita	0.023	0.001
Labour productivity	0.913	0.988
Trade cost	0.000	0.000
Terms of trade	0.060	0.000
Government consumption	0.000	0.000
Investment rate	0.000	0.001
Old-age dependency ratio	0.000	0.000
CPI-based RER	0.000	0.000
GDP deflator-based RER	0.001	0.000
PPP-based RER	0.008	0.000
PPI-based RER	0.000	0.000
ULCT-based RER	0.000	0.002

Notes: H0: All panels are non-stationary; H1: A fraction of the panels are stationary. The US is excluded from the sample.

A3. Cointegration tests

Westerlund's (2005) test

Let y_{it} be an I(1) dependent variable of the following panel-data model, for *i=1,…,N* and *t=1,…T*:

(A21) $y_{it} = \boldsymbol{\beta_i}' \boldsymbol{x_{it}} + \boldsymbol{\gamma_i}' \boldsymbol{z_{it}} + u_{it}$

where $\boldsymbol{x_{it}}$ is a (kx1) vector of I(1) variables, $\boldsymbol{\beta_i}$ denotes the cointegrating vector, which may vary across individuals in the group-mean test employed herein, and $\boldsymbol{z_{it}}$ is a vector of deterministic terms, including individual-specific fixed effects and cross-section means, which thereby account for CSD. Westerlund's (2005) cointegration test is based on the null hypothesis that y_{it} and $\boldsymbol{x_{it}}$ are not cointegrated and tests the stationarity of u_{it}. Rejection of the null hypothesis implies that u_{it} is stationary and that y_{it} and $\boldsymbol{x_{it}}$ are cointegrated in at least some cointegration units.

Westerlund's variance ratio (VR) test statistic is constructed by fitting the model in (A22) using OLS, obtaining the predicted residuals ($\widehat{u_{it}}$) and then testing for a unit root in the predicted residuals using the DF regression model:

(A23) $\widehat{u_{it}}=\rho_i\widehat{u_{it}} + v_{it}.$

The test statistic is given by:

(A24) $VR = \sum_{i=1}^{N}\sum_{t=1}^{T}\widehat{E_{it}^2}\sum_{i=1}^{N}(\widehat{R_i^2})^{-1}$

where $\widehat{E_{it}}= \sum_{j=1}^{N}\widehat{u_{ij}}$ and $\widehat{R_i}= \sum_{t=1}^{T}\widehat{u_{ij}}^2$. It is the group mean of individual variance ratios. The asymptotic distribution of VR, after appropriate standardization, converges to N(0,1).

Table A5 provides Westerlund's (2005) cointegration test results for all ten specifications of the annual BEER model.

Table A5. Testing for cointegration in the annual panel dataset according to Westerlund's (2005) test

	GDP per capita		Labour productivity	
	Variance ratio statistic	P-value	Variance ratio statistic	P-value
CPI-based RER	-2.273	0.012	-2.259	0.012
GDP deflator-based RER	-1.801	0.036	-1.960	0.025
PPP-based RER	1.279	0.100	-0.569	0.285
PPI-based RER	-2.385	0.009	-2.215	0.013
ULCT-based RER	-1.940	0.026	-2.253	0.012

Notes: H0: no cointegration; H1: cointegration in at least some panels. The US is excluded from the sample.

A4. CCEMG estimation

As in Pesaran (2006), Kapetanios et al. (2011) and Eberhardt (2012), assume the following model: for i = 1, . . . , N and t = 1, . . . , T let

(A25) $y_{it} = \beta_i\, x_{it} + u_{it}$

where: $u_{it} = \alpha_{1i} + \lambda_i f_t + \varepsilon_{it}$, $x_{it} = \alpha_{2i} + \lambda_i f_t + \gamma_i g_t + e_{it}$, x_{it} and y_{it} are observables, β_i are country-specific slopes on the observable regressors and ε_{it} and e_{it} are assumed white noise. The unobservables in u_{it} are made up of both fixed effects α_{1i}, which capture time-invariant heterogeneity across countries, and of an unobserved common factor f_t with heterogeneous factor loadings λ_i ,which can account for time-variant heterogeneity and CSD. The CCEMG approach does not require an *a priori* knowledge of the number of unobserved common factors, which may be larger than one (but cannot be larger than the number of variables included in the model); generally the number of unobserved common factors is fixed and small. The factor g_t is included to show that the observables x_{it} are also driven by factors other than f_t. y_{it} , x_{it}, f_t and g_t may be non-stationary. The presence of f_t in both u_{it} and x_{it} raises endogeneity issues, which not tackling the CSD implies. The key passage is that $f_t = \bar{\lambda}^{-1}(\bar{y}_t - \bar{a} - \beta' \bar{x}_t)$ for N→ ∞ since $\bar{\varepsilon}_t$ =0 (iff $\bar{\lambda} \neq 0$). The country-specific equation is thus augmented to include the cross-section averages of the dependent and independent variables, $\bar{y}_t$ and $\bar{x}_t$, which approximate linear combinations of the unobserved common factor f_t:

(A26) $y_{it} = \beta_i\, x_{it} + d_{1,i}\bar{y}_t + d_{2,i}\bar{x}_t + e_{it}$

Because the relationship is estimated for each country separately, the heterogeneous impact (λ_i) is also given by construction. As this is a mean group procedure, the parameters are estimated country-by-country and then averaged across countries:

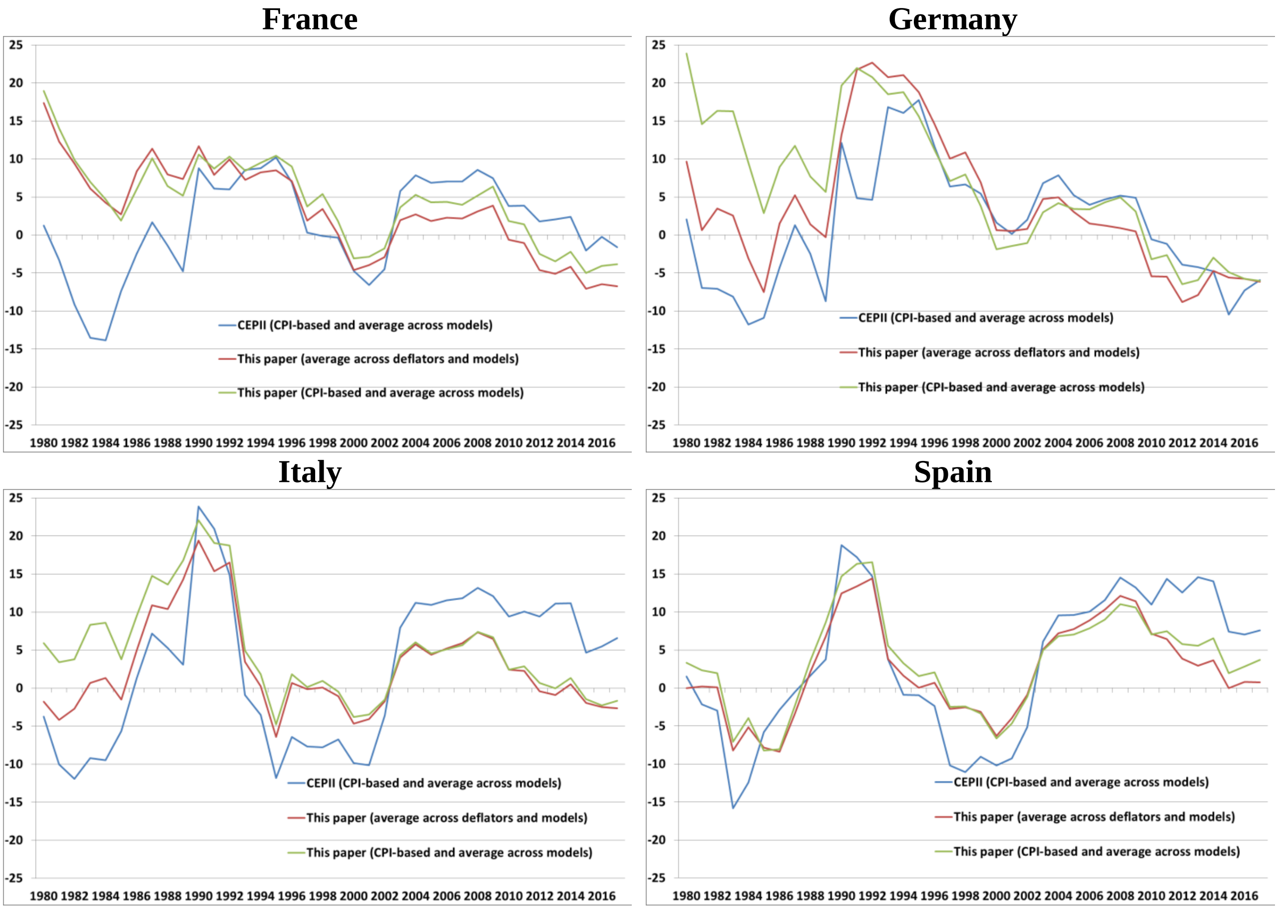
(A27) $\widehat{\beta_{CCEMG}} = N^{-1} \sum_i \widehat{\beta_{CCE,i}}$

Under some general conditions described in Pesaran (2006) and in Kapetanios et al. (2011), $\widehat{\beta_{CCEMG}}$ is an asymptotically unbiased and consistent estimate for β, even if the regressors are weakly exogenous, or if the deviations β_i-E(β_i) are correlated with the regressors/errors. The other estimated parameters have no economic meaning.

Annex B. Additional charts on annual HCI misalignments

Figure B1. Annual HCI misalignments of the main euro-area countries according to two annual BEER models

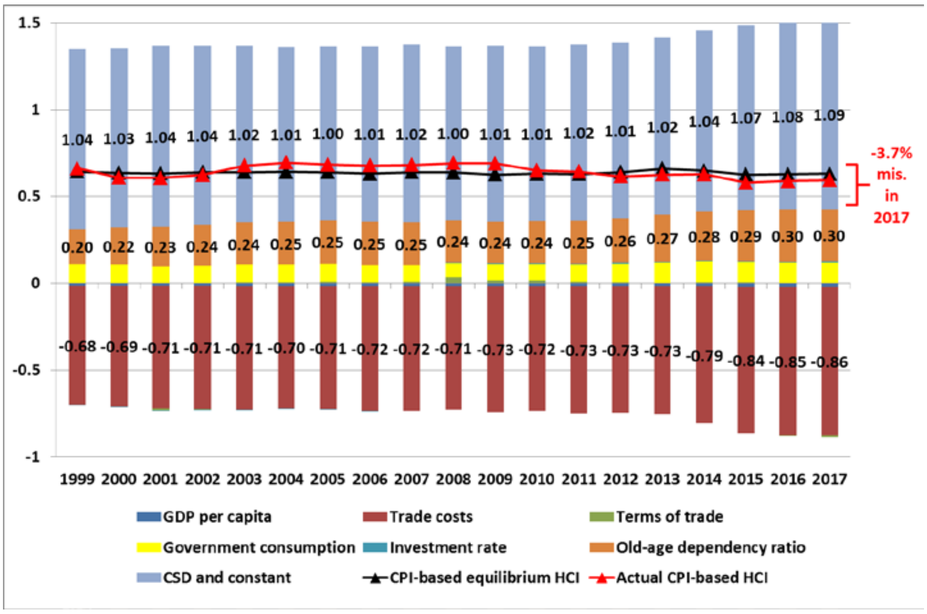
(percentage points)



Source: author's average and CPI-based estimates, based on the annual BEER model described herein, and CEPII CPI-based estimates from the EQCHANGE database (2018 vintage), described in Grekou (2018a).

Figure B2. The "pseudo-contributions" of economic fundamentals to the annual HCI misalignments of the four main euro-area countries, 1999-2017

France



Germany

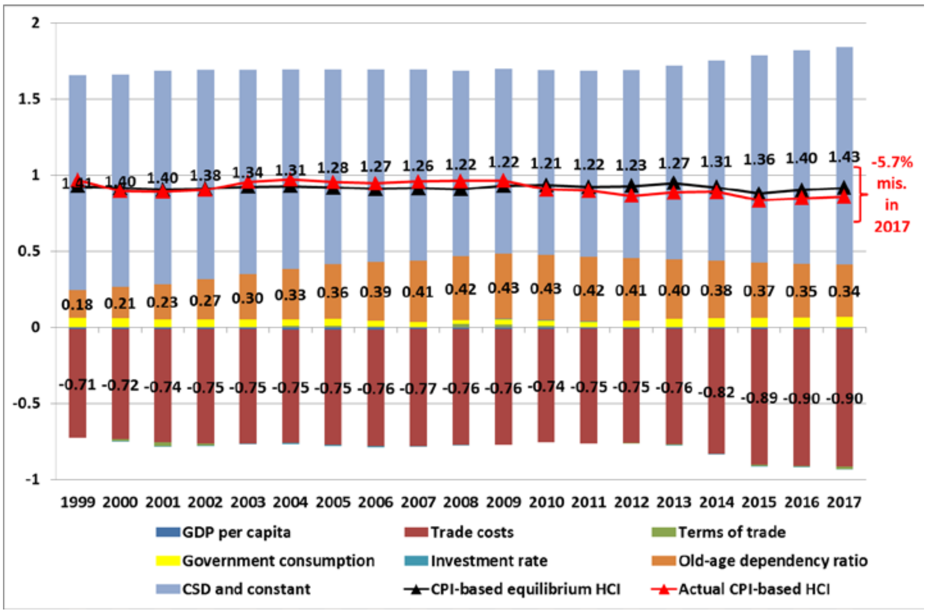
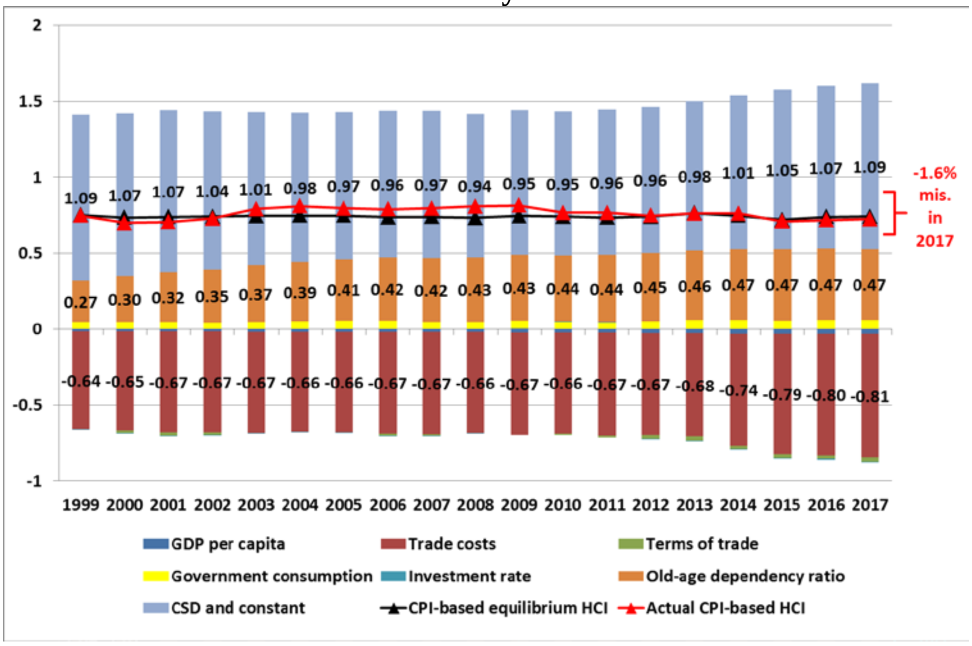
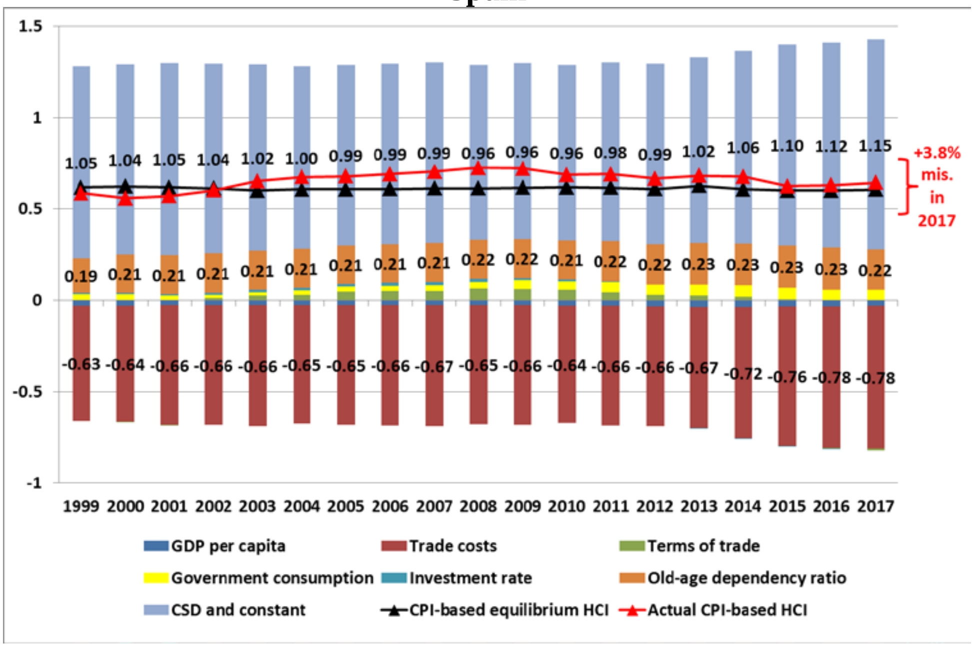


Figure B2 cont.

Italy



Spain



Source: author's estimates.

Notes: The charts refer to the CPI-based estimates, based on the specification including relative GDP per capita as the proxy of the BS effect. "CSD and constant" includes the contribution of the dependent and explanatory variables' cross-section means, which account for cross-sectional dependence.

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