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An exchange market pressure measure for cross country analysis

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ABSTRACT

EMP measures in the existing literature are oriented towards applications in crisis dating and prediction. We propose a modified EMP measure where cross-country comparisons are possible. This is the sum of the observed change in the exchange rate with an estimated counterfactual of the magnitude of the change in the exchange rate associated with the observed currency intervention. We construct a multi-country dataset for EMP in each month. This opens up many new research possibilities.

1. Introduction

The concept of exchange market pressure was first proposed by Girton and Roper (1977). The notion is straight forward: EMP measures the total pressure on an exchange rate, which has been resisted through foreign exchange intervention or relieved through exchange rate change. The problem is that measuring EMP requires combining the observed change in the exchange rate, which is a percentage change, with an observed intervention which is measured in dollars.

The early efforts in measuring EMP worked directly using monetary models, whereas more recent efforts have focussed on measuring EMP using indices that combine changes in reserves and the exchange rate. The direct measures of EMP are model dependent and primarily geared towards finding the magnitude of money market disequilibrium that must be removed either through reserve or exchange rate changes under any desired exchange rate target. EMP indices meanwhile, are designed to capture and forecast crises. Direct measures often lack consistent units, whilst indices do not have this problem and are better suited to crisis conditions.

Girton and Roper (1977) assumed that foreign exchange intervention was unsterilised. Thus intervention led to equivalent amounts of changes in base money. Money was assumed to be neutral so that percentage changes in base money led to equivalent changes in prices. The assumption of purchasing power parity meant that percentage changes in domestic prices were essentially equal to exchange rate changes. Under these assumptions, the authors added the percentage changes in

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reserves and in exchange rates. However, monetary models had low predictive power for changes in exchange rates and the resulting measure of EMP was often misleading (Eichengreen et al., 1996).

When EMP is measured as:

$$E_{GR} = \Delta e_t + \Delta \bar{r}_t$$

the first right-hand side term is in the units of percentage change of the exchange rate, and the second is in the units of percentage change of reserves as a fraction of monetary base. This formula could only been motivated by the assumption that for all countries, at all time periods, a reserves change of 1% of m_0 (monetary base) has an impact on the currency of 1%. But there is no basis for expecting the foreign exchange market to have such a property for all countries and for all times.

In order to address these problems, Eichengreen et al. (1996) created a new measure of EMP. They normalised all prices and quantities, then weighted these components of the index by the inverse of their historical volatilities. Alternative weighing schemes were proposed by Sachs et al. (1996), Kaminsky et al. (1998), Pentecost et al. (2001), Klaassen (2011), and IMF (2007). This approach has led to many useful and important applications in international finance and macroeconomics.

The EMP indices, however, have well documented problems with the "arbitrary" choice of index weights and crisis thresholds (Pontines and Siregar, 2008). In addition, normalisation means the EMP indices cannot be used for cross-country comparisons; they are designed for comparison across time series of a country to indicate periods of "extreme" EMP. Under a fixed exchange rate, many of the conventional measures yield an EMP of infinity, which hampers applications.

Consider a research question such as the impact of quantitative easing (QE) upon emerging markets. It is natural to look at this common shock (QE) inducing exchange market pressure upon all EMs. An array of questions can then be asked. What were the country characteristics which led to high EMP in some emerging markets (EMs) but low EMP in others? Which EMs allowed EMP to be expressed as exchange rate fluctuations, and which EMs did not? What were the causes and consequences of fear of floating? These questions require measurement of EMP in a way that permits comparisons across countries and time.

Consider a practical question such as the outcome of the US presidential election in 2016. It would be useful to observe the EMP and exchange rate changes across all countries of the world in November and December 2016. This could be a useful tool for finance practitioners and for policy makers. This also requires measurement of EMP in a way that permits comparisons across countries and time.

Towards this objective, we build on Weymark (1995), who added the change in the exchange rate that was observed with that of the change in the exchange rate that was *prevented* by the central bank through intervention or by changes in the policy rate. This measure has a consistent unit: the percent change in exchange rate over a one-month period. This takes us to the question: *What is the magnitude of the exchange rate movement associated with $1 billion of intervention?* There are many problems in estimating this. Intervention and the exchange rate level may be endogenous. The impact of foreign exchange intervention, when there is any impact at all, may be asymmetric depending on the direction of the intervention, time varying and temporary (Menkhoff, 2013; Disyatat and Galati, 2007; Lahura and Vega, 2013).

We turn to concepts from the working of financial markets, where the notion of 'market impact' is used when understanding the price change associated with large trades placed by investors. The impact of a billion dollars of intervention depends on the size and liquidity of the currency market.

We propose an estimation strategy through which the exchange rate change associated with $1 billion of intervention is measured for some countries for some points in time. Our method relies on situations where a country has switched between a fixed and a floating exchange rate regime (or vice versa). Assuming similarity of macroeconomic shocks before and after the change in the exchange rate regime, we are able to obtain an estimate of the exchange rate change associated with $1 billion of currency intervention. In our analysis of the global data, we find 39 country-periods where such estimation is possible.

Regression analysis of these values is used to impute values for other country-year settings. This gives the ability to measure EMP for all countries for all months, in a way that is comparable across countries and months. We apply a series of sanity checks and robustness checks, and find that this database has meaningful properties.

This is a paper focused on measurement. It results in a dataset with information about monthly EMP for a large panel of countries. The dataset is released in the public domain and is regularly updated by the authors. An array of interesting research questions, and real world applications, could flow from this work.

2. Measures of EMP

An EMP index consists of a sum of a standardised change in the exchange rate and a standardised change in reserves, both of which are dimensionless and hence conformable for addition. EMP indices were developed for the purpose of analysing currency crises, one country at a time. Crisis periods, in general, are periods when policy makers were often trying to defend the exchange rate, using all possible policy options. All components of EMP are generally seen to move up in this period. When each of these is first standardised, and then added up to obtain an index, the index has high values for periods of crisis, where high is often identified as the index being some standard deviations away from the norm. These indices are, however, not appropriate for cross-country comparisons.

We define I_t as the intervention of the central bank in time t. The exchange rate is denoted by e_t, reserves by r_t, base money as m_0 and reserves divided by base money by $\bar{r}_t$. The change in e_t is denoted by Δe_t; the change in r_t is denoted by Δr_t. The change in $\frac{r_t}{m_0}$ is denoted by $\Delta \bar{r}_t$. Under this notation, some of the existing EMP measures are:

E_{et} (Eichengreen et al., 1996):

$$\mathrm{EMP}_t = \frac{1}{\sigma_e}\frac{\Delta e_t}{e_t} - \frac{1}{\sigma_{\bar{r}}}\left(\frac{\Delta \bar{r}_t}{\bar{r}_t} - \frac{\Delta \bar{r}_{US_t}}{\bar{r}_{US_t}}\right) + \frac{1}{\sigma_i}(\Delta(i_t - i_{US_t})) \tag{1}$$

E_{st} (Sachs et al., 1996):

$$E_{st} = \left(\frac{\frac{1}{\sigma_e}}{K_t}\right)\frac{\Delta e_t}{e_t} - \left(\frac{\frac{1}{\sigma_r}}{K_t}\right)\frac{\Delta r_t}{r_t} + \left(\frac{\frac{1}{\sigma_i}}{K_t}\right)\Delta i_t \tag{2}$$

$$K_t = \frac{1}{\sigma_e} + \frac{1}{\sigma_r} + \frac{1}{\sigma_i} \tag{3}$$

E_{kt} (Kaminsky et al., 1998):

$$E_{kt} = \frac{\Delta e_t}{e_t} - \left(\frac{\sigma_e}{\sigma_r}\frac{\Delta r_t}{r_t}\right) + \left(\frac{\sigma_e}{\sigma_i}\Delta i_t\right) \tag{4}$$

and E_{pt} (Pentecost et al., 2001): an index based on a principal components analysis of the sub-components in Eichengreen et al. (1996).

E_{imf} (IMF, 2007):

$$\mathrm{EMP} = \frac{1}{\sigma_{\Delta\% e_{i,t}}}\Delta\% e_{i,t} + \frac{1}{\sigma_{\Delta\% res_{i,t}}}\Delta\%\mathrm{res}_{i,t} \tag{5}$$

$$\Delta\%\mathrm{res}_{i,t} = \frac{\mathrm{NFA}_{i,t} - \mathrm{NFA}_{i,t-1}}{\text{Monetary base}_{i,t-1}} \tag{6}$$

$$\Delta e_{i,t} = \frac{\mathrm{er}_{i,t} - \mathrm{er}_{i,t-1}}{\mathrm{er}_{i,t-1}} \tag{7}$$

In all these cases, for a fixed exchange rate regime, the standard deviation of the exchange rate which is the term in the denominator, is zero. This results in giving an infinitely large weight to the coefficient of exchange rate movements. Consequently, when a country with a pegged exchange rate allows small changes in the exchange rate to occur, these show up as a high EMP because of the large weight being given to exchange rate changes. As an example, consider a historically inflexible exchange rate like that of China, where for long periods of time $\sigma_{\Delta e} \approx 0$. In periods when a small exchange rate change takes place, and the numerator is non-zero, a very large and spurious value for EMP will be induced.

We see the consequence of the large weight given to exchange rate movements, and the small weight given to intervention due to the large variation in intervention in the EMP measure shown in Fig. 1.[1] The EMP appears low in periods when there was a large change in reserves, and higher when there was a small change in the exchange rate.

Conversely, for floating exchange rate regimes, these measures give a large weight to intervention. In addition, there are measurement issues, as not all countries release intervention data. When reserve changes are used to approximate intervention in countries where the exchange float is relatively clean, revaluation effects and interest income end up being given a large weight due to the low variance of reserves. These show up as large EMP, spuriously signalling heavy exchange market pressure.

These characteristics make conventional EMP indices unsuitable for comparisons across countries. As an illustration Fig. 2 shows the EMP index for 4 countries: China, India, Brazil and Egypt. Each country's EMP depends on its historical experience as the measure uses standard deviation from historical data. As a consequence, there are no visible differences between the EMP for a country like China that witnessed large appreciation pressure during the 2000s and the others that witnessed smaller exchange market pressure in both directions. The usefulness of other EMP indices for cross country comparisons varies, but the essential argument for not using them for such comparisons remains unchanged.

2.1. A new EMP measure

In order to do cross country comparisons as well as comparisons across time, we propose an EMP measure with consistent units - percent change in the exchange rate. The proposed measure adds the change in the exchange rate that took place, and the change that we expect would have occurred had there been no intervention. Both components are measured in the same

[1] We use the EMP measure given in Eq. (1) above, that is, as defined in Eichengreen et al. (1996). This measure will be used for all subsequent references to EMP index.

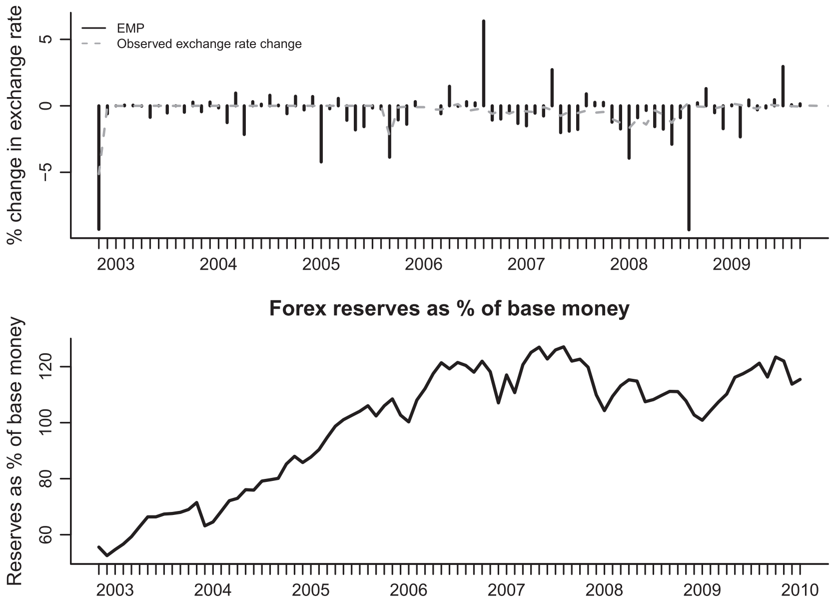

Fig. 1. EMP index for China and the foreign exchange reserves build up. This panel juxtaposes the EMP index calculated for China, with observed exchange rate change and forex reserves as a percentage of base money. The EMP index indicates small/zero appreciation pressure on the renminbi from 2003 to 2006. However, the massive reserve build up during the same period seems to suggest the renminbi was under strong pressure to appreciate.

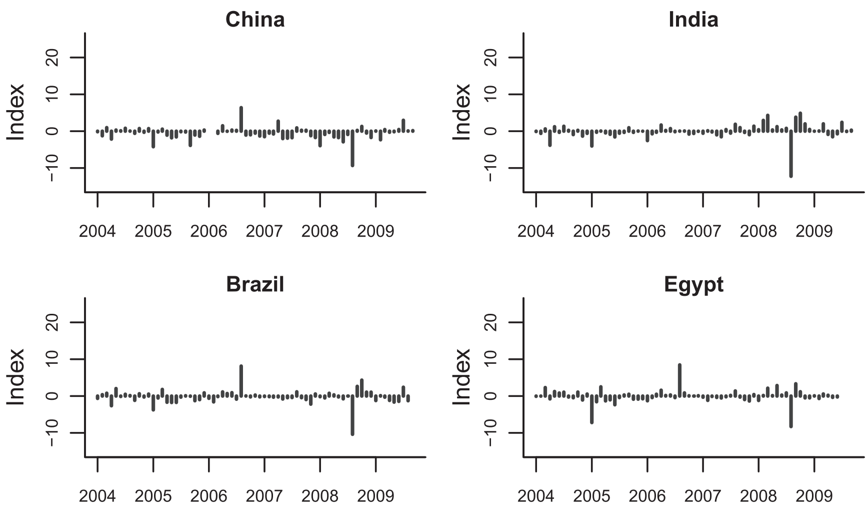

Fig. 2. The EMP index for four countries. This panel shows the EMP index for China, India, Brazil and Egypt. The magnitude of the EMP appears similar across the four countries. Using this measure, it is not possible to tell whether India witnessed a different magnitude of pressure than China in the 2000s.

units, i.e in terms of the percentage change in the exchange rate. To transform the intervention into a measure of the percentage change that was prevented, we need a conversion factor, which we denote as ρ_t. The challenge is that as conditions in foreign exchange markets evolve, the impact of intervention may vary. ρ_t is not a constant and may be expected to vary over time, and across countries.

We propose to measure EMP *in units of percentage exchange rate change over a one-month period:*

$$\mathrm{EMP}_t = \Delta e_t + \rho_t I_t$$

- *Δe_t is the percentage change in the exchange rate,*
- *I_t is the intervention measured in billion dollars,*
- *ρ_t is the conversion factor, which is the change in the exchange rate associated with \$1 billion of intervention. The value of the conversion factor will depend on size and liquidity of the foreign exchange market.*

It follows that $\rho_t I_t$ is expressed in units of percentage change of the exchange rate. It is the exchange rate change of the month we would have expected if there had been no intervention. A key question is whether the conversion factor can be estimated sufficiently well to produce a robust measure of EMP. In this paper, we propose an empirical strategy for measuring ρ_t. We show that these values are consistent with our priors about what ρ_t ought to be. We go on to utilise these values to construct an EMP database, which has attractive properties.

2.2. Estimating ρ_t

Estimates of the impact of intervention in the literature vary highly due to identification problems. The impact depends on other policies such as sterilisation, communication or inflation targeting by the central bank (Menkhoff, 2013). For example, Evans and Lyons (2006) estimate the impact that ordinary order flow has on the exchange rate as 0.44 basis points per 10 million US dollar order flow in the highly liquid Deutsche Mark-US Dollar market in 1996. Scalia (2008) estimates an impact between 7 and 12 basis points per 10 million euro for the Czech Republic. Tapia and Tokman (2004) estimate that in Chile, sales of US dollar in 1998–99 resulted in a 1 percent exchange rate change on 500 million US dollar intervention. Guimares and Karacadag (2004) estimate that 100 million US dollar sales has an impact on the Mexican peso of 0.4 percent, whereas purchases have no effect. Though the estimates are not strictly comparable, ρ estimates in this literature, when translated into our framework, lie between 0 and 10 percent impact upon the exchange rate, of a billion dollars of intervention.

Whilst these papers are useful for obtaining an intuitive sense of the plausible magnitudes, for the purpose of a data-driven algorithm that utilises cross-country data to create a panel database about EMP, we require an estimation strategy which yields estimates of ρ_t on a global scale. We propose going about this in two steps. The first step is to estimate ρ_t in certain situations in the data. The second step is to find the determinants of the estimated ρ_t and to use these to predict ρ_t for all country periods and years.

The first step is based on a key insight which yields an identification opportunity. Assume a country which has experienced both fixed and float periods. Assume during the fixed periods, the country only uses intervention to influence the exchange rate, and that it does no intervention during the float periods. These are highly restrictive assumptions, but necessary to permit identification of ρ (we return to this issue in Section 3). Accordingly, we observe Δe_t in float periods and I_t in fixed periods.

$$\mathrm{EMP}_t = \Delta e_t + \rho_t I_t$$

$$\mathrm{EMP}_{\mathrm{float}} = \Delta e_t$$

$$\mathrm{EMP}_{\mathrm{fixed}} = \rho_t I_t$$

In order to identify "normal times" which do not have unusual macroeconomic volatility, we exclude countries with currency crises. *In normal times, we argue that macroeconomic shocks and hence* EMP *volatility are similar across these periods* and consequently, EMP volatility. Under this assumption:

$$Var(\mathrm{EMP}_{\mathrm{fixed}}) = Var(\mathrm{EMP}_{\mathrm{float}}) \tag{8}$$

$$\rho_t = \left(\frac{Var(\Delta e_{\mathrm{float}})}{Var(I_{\mathrm{fixed}})}\right)^{\frac{1}{2}} \tag{9}$$

This gives an opportunity for measuring ρ_t in some situations. To estimate ρ, we need to observe countries which have experienced both fixed and floating exchange rate regimes. These should be periods in which we can assume that the volatility of the exchange market pressure is roughly constant. The fixed and float regimes should be adjacent so that this is a relatively short window of time.

We analyse 137 countries from February 1995 to December 2009, using the Zeileis et al. (2010) methodology to identify structural breaks in the *de facto* exchange rate regime. This methodology finds dates of structural change in the Frankel and Wei regression (Frankel and Wei, 1994). The R^2 of the Frankel-Wei regression is our measure of exchange rate flexibility. For this purpose, we define a fixed exchange rate regime as a period when $R^2 > 0.95$, and a floating exchange rate regime when $R^2 < 0.66$. Each period is required to be at least 12 months long.

The dates for structural change of the exchange rate regime are validated against the Reinhart and Rogoff (2004) exchange rate regime breaks (Appendix A.1). We exclude periods where macroeconomic shocks were known to be high in one of the periods and known crisis dates. We also remove periods defined as "freely falling" by Reinhart and Rogoff (2004), when the volatility of the exchange market pressure cannot be assumed to be constant. This gives us 26 events where a country moved from a floating to a fixed exchange rate regime, and 13 events where a country moved from a fixed to a floating exchange rate regime.[2]

[2] Appendix A shows that periods identified by us as float roughly match the Reinhart and Rogoff (2004) classification of managed float.

We estimate ρ using the above methodology for each of the 39 regime break points in our dataset for which such assumptions can be made. Table 1 show the estimated values of ρ_t associated with episodes of break dates that involve a movement from a floating exchange rate regime to a fixed exchange rate regime. Table 2 shows the values of ρ_ts estimated when countries move from a fixed exchange rate regime to a floating regime. For every episode, the value of ρ_t is attributed to the mid-point of the window for estimation.

As an illustration, we show all the steps involved in estimating ρ_t for one example: Kenya. This is one of the countries seen in Table 1 which moved from a floating rate to a fixed rate. Fig. 3 shows the dates of structural break of the exchange rate regime. From April 1997 to July 2001, the Kenyan shilling was floating. This is followed by a period from July 2001 till December 2002 when the Kenyan Shilling was pegged to the USD and the Kenyan central bank was intervening in the currency market. Reinhart and Rogoff (2004) identify this entire period as a de facto crawling peg regime: this highlights the improvements in exchange rate regime analysis obtained using the ZSP methodology.

Using Eqs. (8) and (9), our estimation of ρ_t is 105 percent per billion dollars. This suggests that a million dollars of intervention by the Central Bank of Kenya in currency markets would have prevented a 0.105% change in the exchange rate in the period July 2001 to December 2002. The number makes intuitive sense when compared with other estimates of the impact of intervention.

Table 1
Episodes of transition from float to fixed. This table shows 26 structural change events for currency regime from float to fix. These country-periods have been used to estimate ρ. As an example, Angola switched from a float to a fix in May 2007, and this episode yields an estimate of $\rho_t = 3.08$; i.e. a billion dollars of intervention would yield a 3.08% change in the exchange rate.

Country	Float period	R^2	Fix period	R^2	ρ_t
Angola	November 2006–May 2007	0.55	May 2007–February 2009	0.99	3.08
Bangladesh	December 2005–January 2007	0.62	January 2007–October 2010	0.95	6.85
Brazil	June 1994–July 1995	0.51	July 1995–January 1999	0.99	1.97
Belarus	June 2009–April 2010	0.59	April 2010–April 2011	0.97	5.41
Cape Verde	March 1999–September 2001	0.31	September 2001–October 2002	1.00	669.26
Djibouti	June 1996–May 1997	0.34	May 1997–December 1999	0.99	604.14
Djibouti	March 2002–October 2002	0.53	October 2002–July 2004	1.00	376.19
Ethiopia	September 2002–May 2007	0.65	May 2007–January 2009	0.96	8.68
Guinea	August 1998–September 1999	0.55	September 1999–August 2001	1.00	268.97
Guyana	October 1998–July 1999	0.48	July 1999–June 2005	0.99	442.47
India	August 1997–August 1998	0.50	August 1998–March 2004	0.97	1.55
Kenya	April 1997–July 2001	0.54	July 2001–December 2002	0.97	105.64
Comoros	July 2004–May 2006	0.48	May 2006–December 2006	0.96	462.02
Kazakhstan	March 2006–September 2007	0.58	September 2007–May 2011	0.99	2.48
Laos	June 2001–November 2001	0.44	November 2001–October 2003	1.00	390.10
Sri Lanka	June 2000–June 2001	0.48	June 2001–April 2002	0.95	28.20
Mongolia	September 1998–March 2001	0.45	March 2001–December 2001	0.96	184.94
Maldives	May 2005–April 2006	0.46	April 2006–January 2007	0.96	79.04
Malaysia	August 1997–August 1998	0.21	August 1998–July 2005	1.00	5.35
Tunisia	September 1990–September 1991	0.47	September 1991–August 1992	0.99	38.69
Trinidad and Tobago	September 1996–October 1997	0.59	October 1997–June 1999	0.99	19.18
Trinidad and Tobago	May 2008–May 2009	0.58	May 2009–September 2010	0.96	7.52
Ukraine	March 2008–November 2009	0.18	November 2009–December 2011	0.99	6.49
Venezuela	February 2002–September 2003	0.34	September 2003–January 2010	1.00	8.82
Vietnam	September 2000–May 2001	0.66	May 2001–March 2008	1.00	1.07
Antigua and Barbuda	February 1996–August 2002	0.63	August 2002–October 2011	1.00	36.59

Table 2
Episodes of transition from fixed to float. This table shows 13 fixed to float country-periods which have been used to estimate ρ along with the ρ estimates for those periods. The R^2 in the fixed periods are very high and those in float periods much lower. As an example, Costa Rica switched from fixed to float in January 1997, which yields an estimate of $\rho_t = 5.83$; i.e. a billion dollars of intervention would yield a 5.83% change in the exchange rate.

Country	Fix period	R^2	Float period	R^2	ρ_t
Costa Rica	March 1996–January 1997	0.99	January 1997–July 1997	0.41	5.83
Cape Verde	May 2003–July 2004	0.99	July 2004–December 2007	0.44	343.33
Djibouti	December 1995–June 1996	1.00	June 1996–May 1997	0.34	206.08
Gambia	July 1997–December 1998	0.95	December 1998–November 2003	0.50	691.35
Guyana	July 1999–June 2005	0.99	June 2005–December 2005	0.49	269.32
Laos	April 2000–June 2001	1.00	June 2001–November 2001	0.44	519.76
Moldova	April 2000–November 2000	0.95	November 2000–May 2001	0.56	208.57
Mauritius	April 2001–December 2002	0.98	December 2002–May 2004	0.62	117.42
Malaysia	November 1989–December 1993	0.96	December 1993–July 1994	0.44	2.58
Tunisia	September 1991–August 1992	0.99	August 1992–January 1994	0.61	25.64
Ukraine	August 2002–April 2003	1.00	April 2003–February 2004	0.59	15.74
Vietnam	November 1997–September 2000	1.00	September 2000–May 2001	0.66	6.41
C African Republic	June 2001–May 2002	0.99	May 2002–January 2004	0.50	520.58

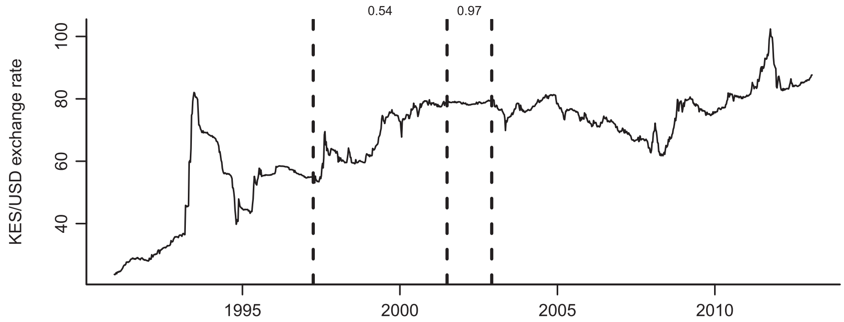

Fig. 3. Exchange rate regimes in Kenya. The graph shows the full history of the Kenyan exchange rate regime. In this, Zeileis et al. (2010) classifies the period from April 1997 to July 2001 as a float with an R^2 of 0.54, and the period from July 2001 to December 2002 as a fixed exchange rate regime with an R^2 of 0.97.

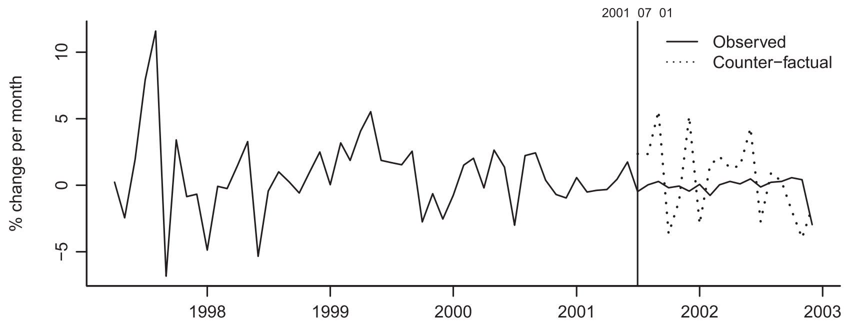

Fig. 4. Kenya: Exchange Market Pressure. The figure presents an estimate of the change prevented in the exchange rate by intervention by the Kenyan central bank when the regime shifted from a float to a fix between July 2001 and December 2002. This suggests that without intervention, we may have observed greater volatility in exchange rate returns during this period.

Fig. 4 shows an estimate of the change in the exchange rate that would have occurred had the central bank of Kenya not intervened in the currency market. This provides a measure of the exchange market pressure in the fixed period. In the floating period, EMP can be seen as the observed change in the exchange rate.

2.3. Predicting ρ_t

The conversion factor ρ_t is primarily about the liquidity of the currency market. The impact of central bank intervention on the foreign exchange market will vary by country, by time. As the size of a currency market changes, ρ will change. We therefore need to estimate a ρ_t time-series for each country to measure EMP. Data for size of the foreign exchange market, in terms of the daily dollar turnover in the spot and derivatives markets is available for some countries and years from the Bank for International Settlement.[3]

The numerical magnitude of ρ will tend to be smaller when the currency market is more liquid, i.e. for bigger and more internationalised countries with greater financial development.[4] Our estimates of ρ for larger emerging markets like Brazil, Turkey, India, Malaysia, Belarus, Indonesia indeed show ρ in the range of 1–10, consistent with the literature (Section 2.2). Meanwhile, countries with very small economies and small foreign exchange markets see a very large impact of a billion dollars of intervention, as in Cape Verde, Guyana and Gambia. In other words, the estimates of ρ – though requiring restrictive assumptions – conform to priors suggested by finance theory.

Table 3 shows the estimated values of ρ_t and the daily turnover in the spot and forwards currency market in or around the same years: Brazil in 1997, India in 2001 and Malaysia in 2002 (Unfortunately, foreign exchange turnover data is not available for most of the country periods for which ρ_t can be estimated). Trading in the foreign exchange market takes place on an average of 20 days a month. In the case of India, for example, the turnover in the market in one month in 2001 was USD 3 billion a day or USD 60 billion per month. Our estimates of ρ suggest that a billion dollars of trade per month by the Indian central bank would have led to a change in the rupee-dollar rate of 1.55 percent in a month in 2001.

[3] BIS Triennial Central Bank Survey of Foreign Exchange and Derivatives Market Activity.

[4] Klaassen and Jager (2011) and BIS (1993) note that the extent of intervention depends on the turnover in the foreign exchange market.

Table 3
Estimated ρ_t and foreign exchange market turnover. This table shows examples of estimated ρ and foreign exchange market turnover. The evidence points towards a negative relationship between currency market turnover and impact of intervention. *Source:* BIS, Brazil data is for 1998 and Malaysia for 2001.

Country	Year	ρ_t	FX market daily turnover (in billion USD)
Brazil	1997	1.97	5
India	2001	1.55	3
Malaysia	2002	5.35	1

The estimation of ρ_t in Section 2.2 gives us values for ρ_t for only 39 country-periods. The missing ρ_ts therefore need to be predicted on the basis of the size of the currency market. Without data on turnover for most country-periods, we proxy it by the size of the economy, financial sector development and integration of the economy with the world economy.

Fig. 5 explores the validity of the proxies in the country-periods for which turnover data and estimates of ρ exist. We expect that as the economy grows bigger, there are more foreign exchange transactions – both the size and the number of transactions would increase. Thus the turnover in the foreign exchange market would be greater. This is seen in the positive relationship between GDP and the turnover in the foreign exchange market. We would also expect that as the size of GDP and the foreign exchange market turnover increase, the impact of a billion dollars of intervention will be lower. The figure shows a negative relationship between ρ_t and GDP. We exploit these relationships to set up a regression model to predict the missing ρ_t.

For prediction of missing ρ_t, since data for the size of the market is not available for all countries and all years, we use the variables that predict foreign exchange market turnover. These include GDP, inflation and various measures of openness of the economy such as the trade to GDP ratio, foreign direct investment and assets and liabilities of the country measured by the Lane and Milesi-Ferretti (2007) measure. Table 4 shows various models for predicting ρ_t. We use model 4 as our base model as the adjusted R-squared does not increase as we add/remove variables in subsequent models. For countries for which financial sector data is not available, missing values of ρ_ts are predicted using only GDP data. Using this method we predict annual values of ρ_t for 172 countries for the years 1995–2011.

2.4. EMP estimates

We now have an annual multi-country dataset of the conversion factor ρ_t required for measuring EMP. For a monthly EMP dataset we assume that the values of the conversion factor remains constant over each year: whilst financial market liquidity fluctuates from day to day, secular changes take place over multi-year time horizons reflecting GDP, internationalisation of

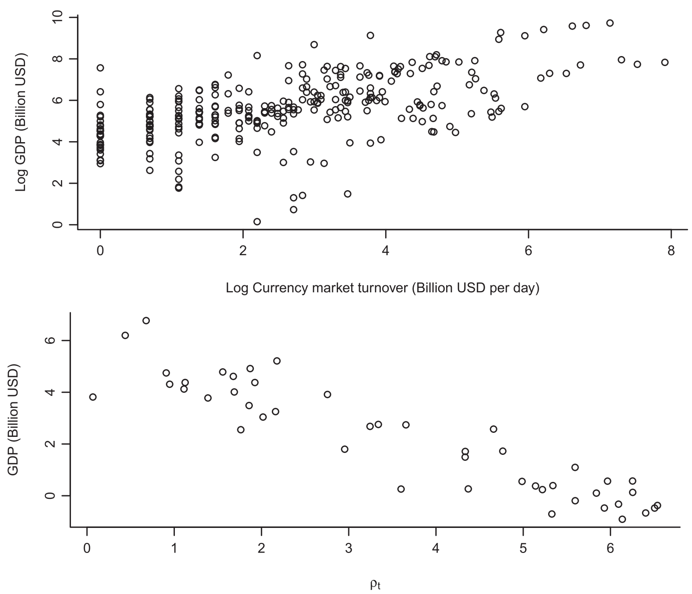

Fig. 5. Relationship between GDP, size of the market and GDP and ρ_t. We expect an inverse relationship between ρ_t and size of the foreign exchange market, or, as the size of the foreign exchange market increases, a billion dollars of intervention by the central bank has a smaller impact. These graphs show that at higher levels of GDP, turnover in the foreign exchange market is higher. Further, at higher levels of GDP, we see that ρ_t is smaller.

Table 4
Model for predicting missing ρ_t. This table displays the various specifications of macro-variables which have been used to model and predict ρ. We use Model 4 for predicting ρ values. Wherever values of trade intensity or FDI to GDP are missing, we use Model 1 with only GDP to predict values of ρ.

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
Intercept	5.55*	8.85*	9.00*	6.22*	6.58*	4.68*	6.72*	6.72*
	(0.17)	(1.02)	(1.11)	(1.55)	(1.68)	(0.51)	(1.71)	(1.71)
GDP	−0.89*	−0.93*	−0.93*	−0.89*	−0.90*	−0.88*	−0.83*	−0.83*
	(0.06)	(0.05)	(0.06)	(0.06)	(0.07)	(0.06)	(0.08)	(0.08)
Trade to GDP		−0.72*	−0.67*	−0.36	−0.35		−0.55	−0.55
		(0.22)	(0.23)	(0.27)	(0.29)		(0.32)	(0.32)
Inflation			−0.21		−0.24	−0.20		
			(0.18)		(0.20)	(0.19)		
Net FDI to GDP				−0.25*	−0.26	−0.33*	−0.30*	−0.30*
				(0.11)	(0.14)	(0.11)	(0.12)	(0.12)
LMF_n							0.47	
							(0.42)	
$LMF^2{}_n$								0.24
								(0.21)
N	46	44	37	38	33	35	31	31
R^2	0.85	0.88	0.88	0.88	0.88	0.88	0.88	0.88
Adj. R^2	0.85	0.87	0.87	0.87	0.86	0.86	0.86	0.86
Resid. sd	0.80	0.74	0.75	0.75	0.76	0.75	0.76	0.76

LMF_n is the Lane-Milesi-Ferreti index after subtracting official reserves.
Standard errors in parentheses.
* Indicates significance at $p < 0.05$.

the economy, and financial sector development. Values of ρ for the latest years for which they could not be predicted due to unavailability of data are assumed to remain unchanged for the last observed year. These ρ estimates are used to compute monthly EMP for all countries in the database (excluding eurozone countries) for the period January 1995–April 2015.

As an example, Fig. 6 juxtaposes our proposed EMP measure for China with the EMP index. In the pre-crisis years of the 2000s, China either witnessed reserve accumulation or the currency appreciated. Consequently, the direction of EMP should be only one way. This is seen in our proposed measure, where values less than zero represent a pressure to appreciate. Our proposed measure captures the pressure on the renminbi to appreciate through the 2000s. In contrast, an attempt to calculate the conventional EMP index for China gives rise to values of infinity for 1999–2002 as the Chinese renminbi was

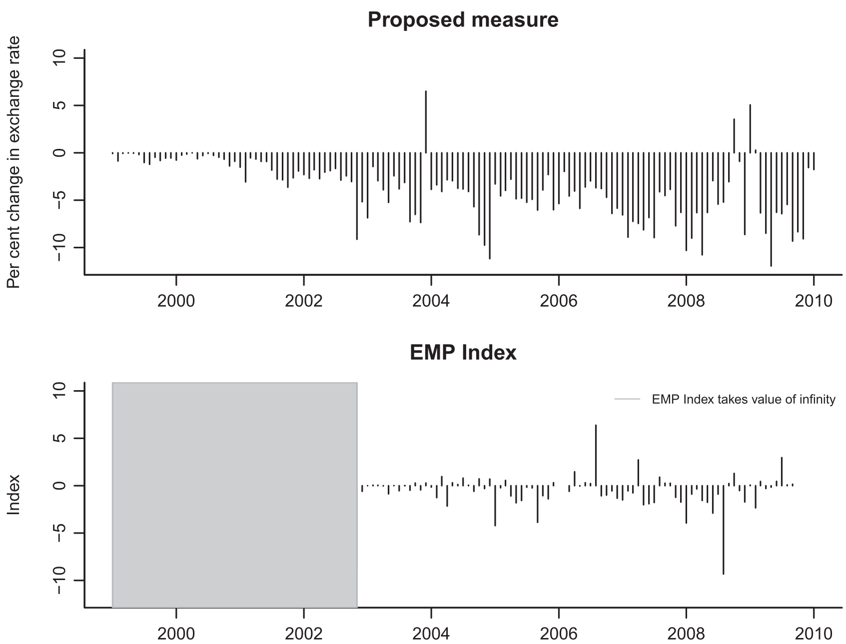

Fig. 6. Comparison between proposed EMP measure and EMP Index for China. This panel shows a comparison between our proposed EMP measure and EMP Index calculated for China for the time period 2000–2010. Our proposed EMP measure suggests that the renminbi has faced increasing pressure to appreciate in pre-crisis years of 2000s. Apart from 4 months when the renminbi depreciated, the renminbi either appreciated or was prevented from appreciating by intervention in foreign exchange markets. The EMP index takes the value of infinity for the period 1999–2002 as $\sigma_{\Delta e} = 0$, whilst our proposed measure works sensibly all through.

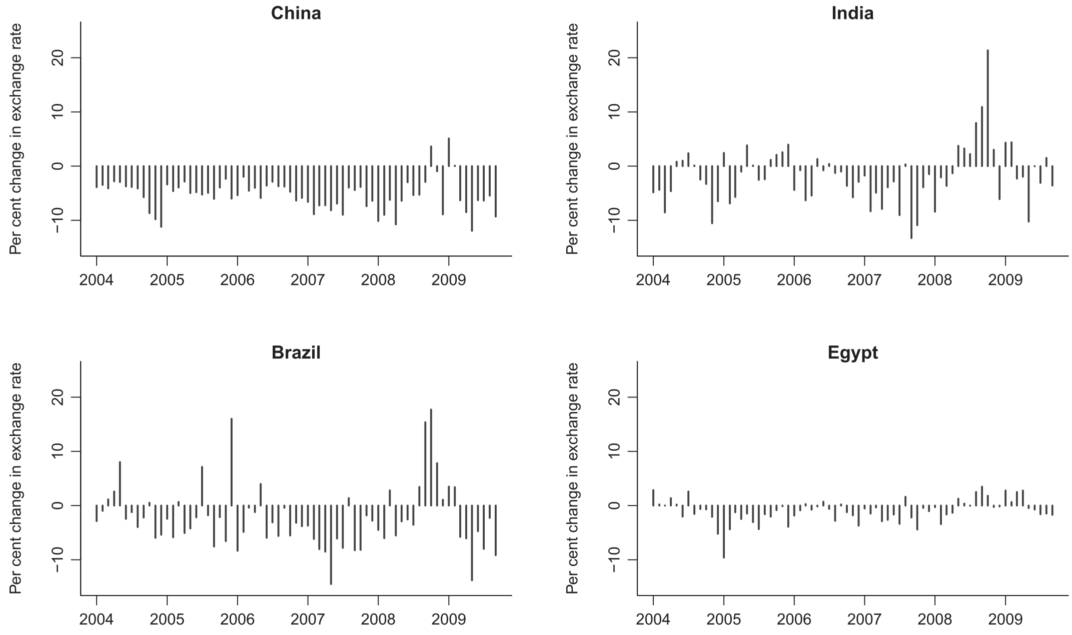

Fig. 7. Proposed EMP measure for selected countries. This panel shows our proposed EMP measure for China, India, Brazil and Egypt calculated for the same time period as in Fig. 1. The figures show a consistent appreciation pressure on the currencies prior to the GFC, consistent with the direction of capital flows during this period.

maintained at a fixed peg of 8.71 per dollar with no variation during this period. The fixed exchange rate with no variation makes the first term in calculation of EMP index ($1/\sigma_{\Delta e} = \infty$) take the value of infinity. Additionally, the uni-directional pressure on the renminbi to appreciate is not evident in the pre-crisis years of 2000s in the EMP index.

Fig. 7 plots our proposed EMP measure calculated for China, India, Brazil and Egypt for the same period as Fig. 2, which plots the EMP Index. India and Brazil witnessed pressure to appreciate for most of the 2000s, with the direction of pressure changing after the 2008 global crisis. Our proposed measure shows this contrasting magnitude and direction of pressures faced by the four countries in 2000s, whilst the EMP index in Fig. 2 seems to indicate that the experience of the four countries has been indistinguishable.

Fig. 8 shows the EMP for Egypt. We see a high pressure on the Egyptian Pound to depreciate with the onset of the Arab Spring. We observe that in the pre-crisis years of 2000's, the Egyptian pound witnessed a sustained pressure to appreciate.

3. Questions on validity of assumptions and estimates

We now examine the various threats to validity. A number of questions may be raised about the data used in the prediction of EMP. Whilst the change in exchange rates (Δe_t) is directly observed, other variables such as conversion factor ρ_t and intervention I_t have been estimated. For the conversion factor ρ_t, we first estimated ρ_t for a small set of countries and then

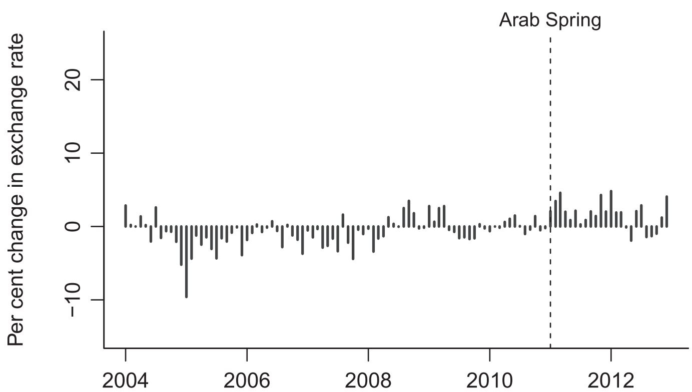

Fig. 8. Proposed EMP measure for Egypt. This panel shows our proposed EMP measure calculated for Egypt for the time period between 2004 and 2013. In the pre-crisis years of the 2000s, we see a consistent pressure on the Egyptian pound to appreciate. After Arab spring, we observe a high pressure on the Egyptian pound to depreciate.

predicted ρ_t for all countries, across time periods, using their determinants, such as GDP. The variable for Central Bank intervention I_t has been estimated by change in reserves. In this section, we address the following questions regarding the validity of our assumptions and accuracy of our estimates:

1. ρ_t estimation: How sensitive are ρ_t estimates to macroeconomic shocks?
2. ρ_t prediction: How good are the predicted ρ_ts?
3. Intervention I_t estimation: How close are the EMP measures in case of countries which publish monthly intervention data?

3.1. Is the estimation of ρ_t's sensitive to macroeconomic shocks?

Though we have dropped country periods for crisis years and freely falling years, it is possible that countries may be moving from fixed to floating because of macroeconomic shocks. If so, this would imply that the currency volatility in the two sets of floating periods, one that precedes and one that follows a fixed regime, would be different. We test whether such a difference exists using the Welch two-sample *t*-test and the two-sample Kolmogorov-Smirnov tests but we find no significant difference in either the means or the distributions. The values for the tests comparing the means and the distribution of the volatility of the exchange rate during the floating period are as follows: The *t*-test gave us a *t*-value of -1.66 with a *p*-value of 0.1 with 40 degrees of freedom. The value of the Kolmogorov-Smirnov statistic was 0.26 with a *p*-value of 0.35. Therefore, we consider both fixed to float and float to fixed episodes in estimating ρ_t.

In the calculations for ρ_t we assumed that the two adjacent periods under consideration had similar macroeconomic volatility. If this assumption is true, then we would expect that macroeconomic shocks should not explain ρ_t. If we regress the calculated ρ_ts on various measures of macroeconomic shocks, the coefficients of these shocks should not be significant. Table 5 shows that ρ_t is not sensitive to variables such as inflation and the current account. We control for GDP, trade integration and capital flows which influence the size of the market and determine ρ_t. None of the coefficients are significantly different from zero. This suggests that the 39 regime changes that were used for estimation of ρ_t were in periods that were not periods of crisis or macroeconomic instability.

Table 5
The assumption of macroeconomic stability. We test the sensitivity of the estimated ρ_t to various measures of macroeconomic shocks across different specifications of a model explaining the ρs. The coefficients for macro-shocks are not significant and this suggests that assumption of macro-stability across our set of corresponding currency regimes holds.

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7
Intercept	5.00*	3.39*	3.26*	6.22*	6.58*	6.31*	5.95*
	(0.90)	(0.32)	(0.31)	(1.55)	(1.68)	(1.92)	(0.45)
Inflation	−0.84				−0.24		−0.23
	(0.47)				(0.20)		(0.21)
CA balance		−0.00					
		(0.04)					
CAD to GDP			−0.06*			0.01	0.01
			(0.03)			(0.02)	(0.02)
Trade int				−0.36	−0.35	−0.37	
				(0.27)	(0.29)	(0.32)	
GDP				−0.89*	−0.90*	−0.92*	−0.90*
				(0.06)	(0.07)	(0.07)	(0.07)
FDI to GDP				−0.25*	−0.26	−0.27	
				(0.11)	(0.14)	(0.14)	
N	39	41	41	38	33	36	36
R^2	0.08	0.00	0.11	0.88	0.88	0.88	0.84
Adj. R^2	0.05	−0.03	0.09	0.87	0.86	0.86	0.83
Resid. sd	2.03	2.04	1.92	0.75	0.76	0.75	0.83

Standard errors in parentheses.
* Indicates significance at $p < 0.05$.

Table 6
Comparing selected estimated ρ_t and predicted ρ_t. Table compares estimated values of ρ with the predicted values of ρ. The predicted values appear comparable and in line with the ρ estimates.

Country	India	Malaysia	Turkey	Brazil	Vietnam	Kenya	Sri Lanka
Year	2001	2001	2002	1997	2001	2002	2001
Estimated ρ_t	1.55	5.35	4.42	1.97	6.42	105.60	28.20
Predicted ρ_t	1.90	4.64	3.64	1.14	9.13	54.97	27.09

3.2. How good are the predicted ρ_t's?

To examine the goodness of our prediction strategy, we now compare the predicted measures of ρ_t using the above model to those originally estimated using the volatilities of the exchange rate and intervention. The comparison can be made only for 38 country-years for which ρ_t could be estimated. Table 6 shows that the predicted ρ values are in the same order of magnitude as the estimates.

Fig. 9 shows that the predicted values of ρ against available estimates are quite close. Fig. 10 shows the correlation between the estimated and the model predicted ρ_t. The two are close to being on a 45° line.

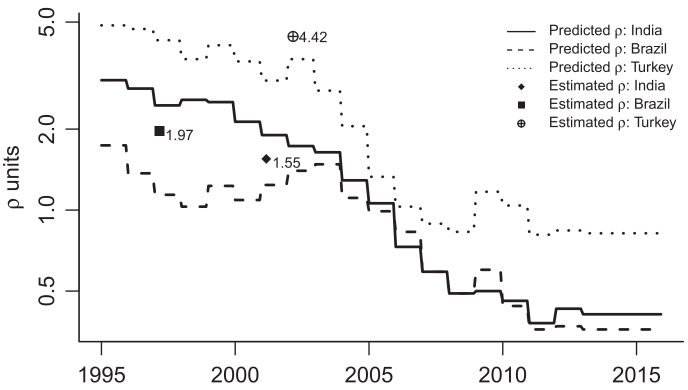

Fig. 9. Selected countries: Predicted and estimated values of ρ_t. Figure compares estimated values of ρ with the model predicted values of ρ for India, Brazil and Turkey. The model predicted values are comparable and in line with the ρ estimates.

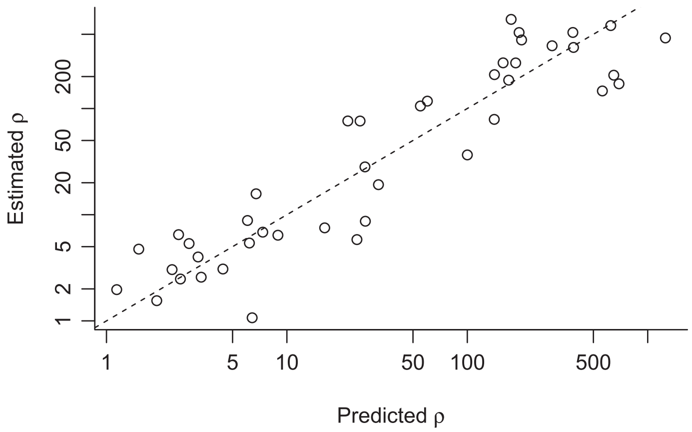

Fig. 10. Scatter plot of model predicted ρ_t versus estimated ρ_t. This figure shows all estimated values of ρ with the corresponding model predicted values of ρ. The predicted values are correlated with the ρ estimates and are close to the 45° line.

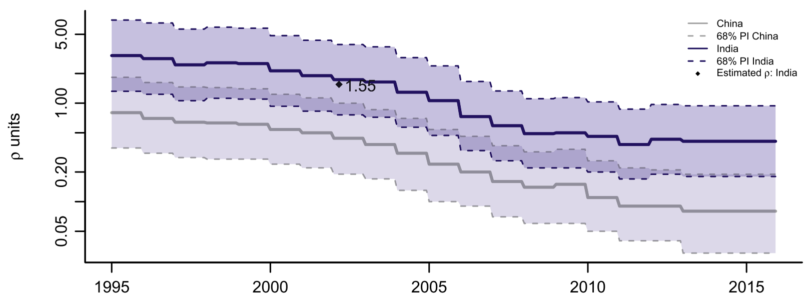

Fig. 11. Predicted ρ_t with $\pm\sigma$ prediction intervals. This figure shows predicted values of ρ with $\pm\sigma$ prediction intervals. The fitted values of ρ are close to the estimated values of ρ all cases and lie within the 68% prediction interval.

We test the stability of the predicted values of ρ by using prediction intervals to get a sense of the probability space of the true ρ parameter. We estimate $\pm\sigma$ prediction intervals for ρ and calculate upper and lower bounds for the ρ estimate (Fig. 11).

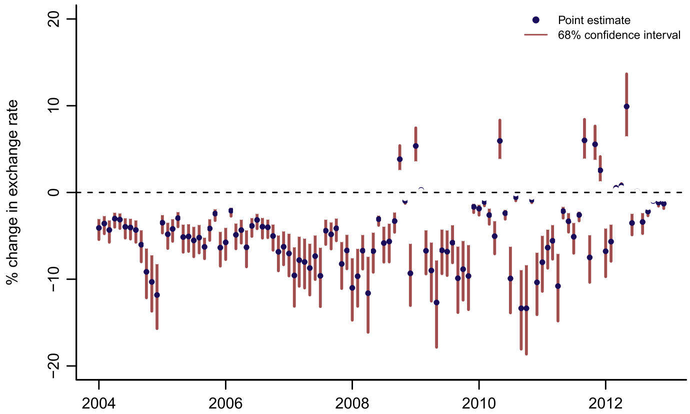

Fig. 12. EMP index for China with confidence interval. This figure plots the EMP index for China. The dots represent the point estimate and the lines represent the 68% confidence interval.

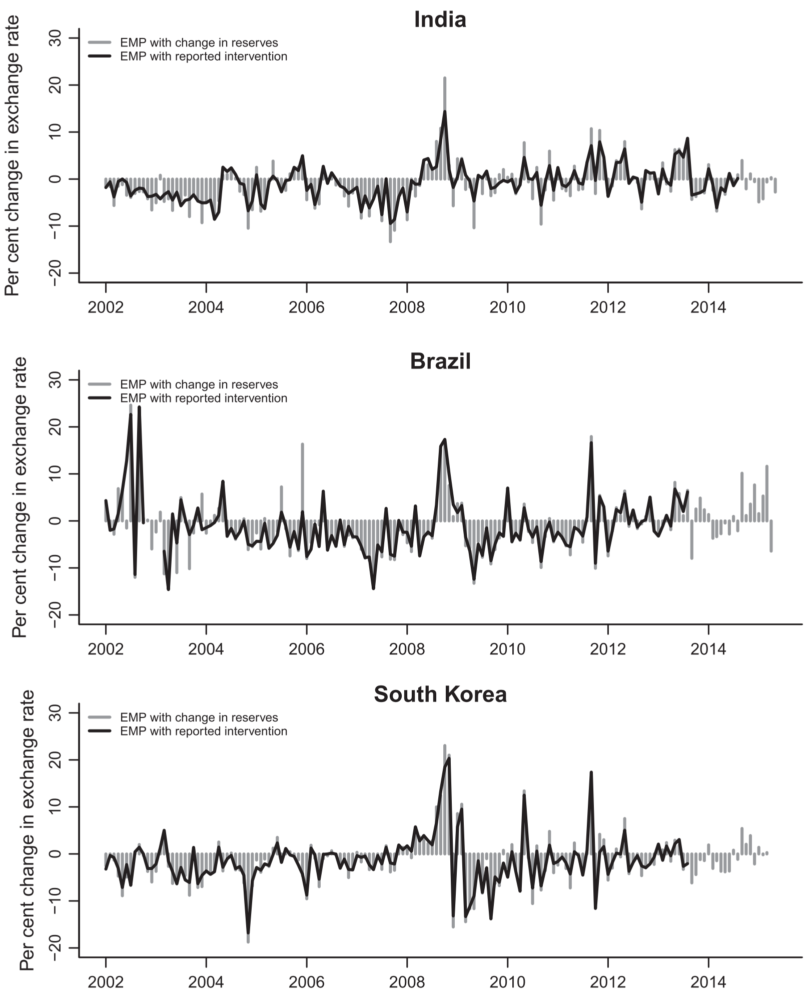

Fig. 13. Intervention data for EMP calculations: Actual versus change in reserves. This figure compares estimates of EMP calculated with reported intervention and change in reserves for India, Brazil and Korea. This suggests that change in reserves are a good proxy for intervention data.

Predicted ρ values are being used to estimate EMP. This necessarily introduces statistical imprecision in the resulting EMP values. We setup a simulation where many draws from the distribution of ρ are utilised to obtain corresponding draws from the distribution of EMP. Fig. 12 superposes the 68% confidence interval with the EMP estimate for China. This shows that whilst the estimate for each month has a wide confidence interval, the overall picture is still useful. In the public release of the data-set, we also release these confidence intervals for ρ and for EMP.

3.3. How robust is the measure of foreign exchange intervention?

Intervention is not reported by most central banks. Consequently the literature uses the change in reserves as a proxy for intervention.[5] But changes in reserves may also happen due to interest payments, or due to revaluation effects. It is not possible to accurately adjust for these without knowing the exact composition of reserves or the timing of interest payments. Further, intervention may be done through swaps, credit lines or intervention in derivatives markets which may not immediately affect reserve levels, but this data is usually not publicly available.

In Fig. 13, we show that when actual intervention data is used for the countries for which central banks release data, the estimates of EMP do not differ significantly from the measures obtained by using the change in reserves.We find central bank intervention time series for 6 countries; India, Brazil, South Korea, Mexico, Russia and Peru. We estimate EMP for these countries using this intervention data and by using the change in reserves data and compare the two measures. The two measures appear similar. This also corroborates recent work by Suardi and Chang (2012) who suggest that changes in reserves are a reliable proxy for central bank intervention.

4. Reproducible research

A database with monthly data for the proposed EMP measure, along with the computer programs used in this research, have been placed on the web.[6] The authors hope this will enable replication and downstream research. The data is available for 139 countries and spans from 1995 to 2012 for most countries (due to limited data availability for some countries). The authors propose to update this database four times a year, and thus make it a useful resource for researchers.

5. Conclusion

Previous EMP measures were employed largely to predict crises. They gave misleading results in more tranquil periods, and could not be used for cross-country analysis.

In this paper we develop a new EMP measure that can be used in normal times, and permits panel data analysis. Since exchange rate changes and intervention are in different units, the paper focuses on creating a conversion factor that allows both to be measured in terms of exchange rate changes, i.e. the change that occurred and the change that was prevented by intervention.

Such a counterfactual can, of course, not be measured accurately. We provide an estimate of the exchange rate change that was prevented, based on a series of restrictive assumptions, most notably that intervention has systematic and durable effects on exchange rate levels, which are related to the size of the market.

The dataset has been released in the public domain and opens up many new academic and policy research possibilities.

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Appendix A

A.1. Defining exchange rate regimes

Estimates of the conversion factor depends upon the transitions from pegs to float and float to pegs. We use Zeileis et al. (2010) (ZSP) to identify these periods. In this appendix, we show that the mapping used from the much more familiar Reinhart and Rogoff (2004) (RR) classification to the R^2 calculated for different country periods by ZSP.

See Tables 7–9.

[5] See Pentecost et al. (2001), Sachs et al. (1996), Kaminsky et al. (1998) and Eichengreen et al. (1996).

[6] http://macrofinance.nipfp.org.in/releases/exchange_market_pressure.html.

Table 7
Comparing RR and ZSP across both datasets. We compare the R^2 calculated for different country periods using ZSP methodology for 137 countries with the RR coarse classification. We compare the RR score with the ZSP R^2 of the Frenkel-Wei regression to ascertain the R^2 thresholds between different de facto currency regimes.

RR score	RR classification	Average ZSP R^2	Max R^2	Min R^2
1	Peg	0.85	1	0.06
2	Crawling pegs	0.81	1	0.16
3	Managed floats	0.61	1	0.08
4	Free floats	0.54	1	0.03
5	Freely falling	0.57	1	0.03
6	Multiple arrangements	0.88	1	0.44

Table 8
Reinhart and Rogoff (2004) monthly-coarse classification. This table describes the Reinhart-Rogoff monthly-coarse currency classification.

Code	Description
1	No separate legal tender
1	Pre announced peg or currency board arrangement
1	Pre announced horizontal band that is narrower than or equal to +/−2%
1	De facto peg
2	Pre announced crawling peg
2	Pre announced crawling band that is narrower than or equal to +/−2%
2	De factor crawling peg
2	De facto crawling band that is narrower than or equal to +/−2%
3	Pre announced crawling band that is wider than or equal to +/−2%
3	De facto crawling band that is narrower than or equal to +/−5%
3	Moving band that is narrower than or equal to +/−2% (i.e., allows for both appreciation and depreciation over time)
3	Managed floating
4	Freely floating
5	Freely falling
6	Dual market in which parallel market data is missing

Table 9
Comparing RR and ZSP for float periods used in the paper. The table shows float periods detected by the Zeileis et al. (2010) (ZSP) methodology and compares it with the Reinhart and Rogoff (2004) (RR) *de facto* coarse currency classification. Majority of the country periods which are detected as floats by the ZSP methodology are categorised as crawling pegs or managed floats by RR database.

Country	Float period	ZSP R^2	RR classification
Angola	November 2006–May 2007	0.55	1
Bangladesh	December 2005–January 2007	0.62	2
Brazil	June 1994–July 1995	0.51	2
Cape Verde	March 1999–September 2001	0.31	2
Ethiopia	September 2002–May 2007	0.65	2
Guinea	August 1998–September 1999	0.55	2
Guyana	October 1998–July 1999	0.48	2
Guyana	June 2005–December 2005	0.49	2
India	August 1997–August 1998	0.5	2
Kenya	April 1997–July 2001	0.54	2
Kazakhstan	March 2006–September 2007	0.58	2
Laos	June 2001–November 2001	0.44	6
Sri Lanka	June 2000–June 2001	0.48	3
Mongolia	September 1998–March 2001	0.45	1
Maldives	May 2005–April 2006	0.46	1
Malaysia	August 1997–August 1998	0.21	4
Tunisia	September 1990–September 1991	0.47	2
Trinidad & Tobago	September 1996–October 1997	0.59	2
Venezuela	February 2002–September 2003	0.34	4
Antigua & Barbuda	February 1996–August 2002	0.63	1
Angola	November 2006–May 2007	0.55	1
Costa Rica	January 1997–July 1997	0.41	2
Cape Verde	July 2004–December 2007	0.44	2
Gambia	December 1998–November 2003	0.5	2
Guyana	June 2005–December 2005	0.49	2
Guyana	January 2007–July 2007	0.47	2
Moldova	November 2000–May 2001	0.56	2
Mauritius	December 2002–April 2004	0.62	2
Malaysia	December 1993–July 1994	0.44	2
Tunisia	August 1992–January 1994	0.61	2
Ukraine	April 2003–February 2004	0.59	1
Central African Republic	May 2002–January 2004	0.5	1

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