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7. Apêndices

Os três ensaios que compõem esta tese foram desenvolvidos ao longo do curso de doutorado. Por esta razão os apêndices que iremos apresentar são na verdade os próprios artigos, que foram escritos originalmente em inglês. Neles são apresentadas todas as derivações dos modelos mencionados ao longo desta dissertação bom como efetuados diversos testes de robustez dos resultados empíricos.

O primeiro apêndice-artigo a seguir, intitulado "*Monetary Policy Credibility and Inflation Risk Premium: A model with application to Brazilian Data*" se refere ao capítulo 2 desta tese. O segundo, intitulado "*Current Account as a Dynamic Portfolio Choice Problem*" *se* refere ao capítulo 3. E, por fim, o terceiro artigo apresentado no Apêndice, "*Cousin Risks: The Extent and The Causes of Positive Correlation Between Country and Currency Risk*" se refere ao quarto capítulo desta tese.

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Monetary Policy Credibility and Inflation Risk Premium: a model with application to Brazilian data*

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Abstract

We derive a simple asset pricing-Taylor rule model to explain Inflation Risk Premium and show that monetary policy credibility is one of its main determinants. We then investigate how credible has been monetary policy in Brazil since 2001. Short run inflation surprises had a significant effect in medium run inflation expectations for most of our sample. This phenomenon leads to a less effective monetary policy, as its output cost is higher. This can be a symptom of at least one of two problems: (i) Inflation inertia due to indexation of the economy; and/or (ii) lack of credibility of the monetary authority. The remedy depends on the cause, for instance, if the reason is simply indexation, central bank independence would not help. As our model suggests, looking at comovements of inflation risk premium and inflation surprises helps to identify if lack of credibility is one of the causes. By doing so, we confirm that this was the case in Brazil untill very recently.

JEL classification: E58, E44, G12, E65

1 Introduction

When monetary policy lacks credibility, central banks have to impose a higher interest rate than otherwise in order to control inflation. Knowing that, many countries have adopted inflation targeting (IT) regime to better anchor inflation expectations and, so, reduce the output cost of monetary policy. Following the spread of IT as the most fashionable monetary arrangement across the globe, the academic literature on monetary policy credibility have flourished. Empirically, the extent of the credibility has been evaluated

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in basically two ways[1]: (i) by running regressions on inflation expectations, as in Cerisola and Gelos (2005) or (ii) by looking at the inflation implicit in nominal and real bonds, the "break-even inflation" as in Svensson (1993), Laxton and N'Diaye (2002) and Gürkaynak et al (2005).

The information contained in this data can be better explored. As reckoned by Bernanke (2004), policymakers have a lot to learn from asset prices but "some of the potentially most valuable information in financial markets often requires considerable theoretical and empirical sophistication to extract". That's what we try to do in this paper. The difference between the yield on nominal and real bonds, referred to as "break even inflation " or "market inflation", is not explained solely by the expected inflation: it include a risk premium as well[2]. Differently from the aforementioned papers, to investigate the credibility of monetary policy we separate the "break-even inflation" into two components: the expected inflation and the inflation risk premium. We present an asset pricing model with a Taylor rule and an equation for inflation dynamics which allow us to interpret this "inflation risk premium" and conclude that monetary policy credibility is one of its main determinants. We apply its results to investigate monetary policy credibility in Brazil.

Other branch of the (very recent) literature is also somehow related to our work as they explore the dynamics of real bonds. Ang and Bekaert (2005) develop a term structure model to separate the nominal yield curve components into expected inflation, real rates and inflation risk premium. It is a reduced form model and in the estimation, they use only data on nominal yield curve and current inflation. Similarly, Buraschi and Jiltsov (2005) present a monetary structural RBC type model and estimate it with Nominal yield and current inflation data. Our approach is different: we use the data on inflation risk premium directly - so there is no need to use any model to decompose it - and our model is derived in order to give us a clear economic interpretation on the determinants of this inflation risk premium data. Our main focus is its application to the study of monetary policy credibility.

A very rich data-set, containing real bonds, nominal bonds and survey of inflation expectation for many horizons allows us to use the methodology proposed to analyze the recent Brazilian experience. An extra reason for choosing this country is that one of the most popular subjects among macroeconomist these days in Brazil is the inflation resilience to the high interest rate. Many possible reasons have been raised: fiscal dominance, subsidized credit[3] with interest rate not sensitive to the short rate determined by the monetary authority, fiscal policy not being tight enough, the short run expansionary effect of the

[1] Some other methods: Rossi and Rebucci (2004) that studies the credibility of the disinflation program in Turkey using a Bayesian approach and Agénor and Taylor (1993) that use the difference between exchange rate quoted in official and black market.

[2] Other component is the liquidity premium since Real bonds are attractive to buy-and-hold investors (such as pension funds), in contrast to nominal securities, which are extensively used for trading and hedging. We ignore this problem in our analysis.

[3] National Development Bank's (BNDES) loans constitute approxematelly 40% of the total availble credit to private sector.

expansion of the credit to consumers and so on. In this paper we will point out that credibility of the monetary authority has been also an important factor hampering the transmission channels of monetary policy. However, recently, there was a substantial improvement in our measure of credibility.

The paper has 5 sections, including this introduction. Section 2 present the empirical evidence that short run inflation surprises affects positively 12 month inflation expectations in Brazil. This could be happening either by inflation inertia or by lack of credibility on monetary authority. How could we know if credibility is part of the explanation for this problem? In section 3 we derive a theoretical model that shows that movements in inflation risk preimium can help to identify that. Section 4 then shows that inflation risk premium is very sensitive to inflation surprises, indicating lack of credibility of monetary policy during most of the sample period analyzed. Section 5 concludes and provides some conjectures on possible causes for the lack of credibility.

2 Evidence of the Effects of Short Run Surprises in Medium Run Inflation Expectations

Before deriving the model of inflation risk premium, we start by motivating the exercise. Why would someone be interested in looking at this risk premium? As we will show later, the answer is: because there is a lot of information contained in it, including information about the expected response of the monetary authority to future inflation shocks. In other words, by looking at that inflation risk premium one could infer if market participants expect a more dovish or more hawkish response by monetary authority in the future. For this reason, in the present section we use the case of Brazil to motivate the model and start the study of monetary policy credibility.

Inflation target regime in Brazil begun in 1999. The Central Bank target is over the consumer price index - IPCA. The targets are usually defined and announced in June[4]. The table below present the targets, objectives and results of inflation targeting in Brazil since the adoption of IT:

Since the targets are valid only for the calendar year, we calculated the 12 month ahead target by interpolating the official calendar-year targets. Through out the paper we work with monthly data, so on the coming notation, each t represents a month. Define "Expected Deviation from the Target" as the expected 12 month ahead inflation minus the interpolated 12 month ahead target:

- Expected Deviation from the Target $= E_t \left(\sum_{s=t+1}^{t+12} CPI_s \right) -$ (Interpolated CB Target for the next 12 months).

Twelve months ahead inflation expectations were collected from Brazilian Central Bank's expectation survey. We present in the Chart 1 below the interpolated Central Bank announced target for the next 12

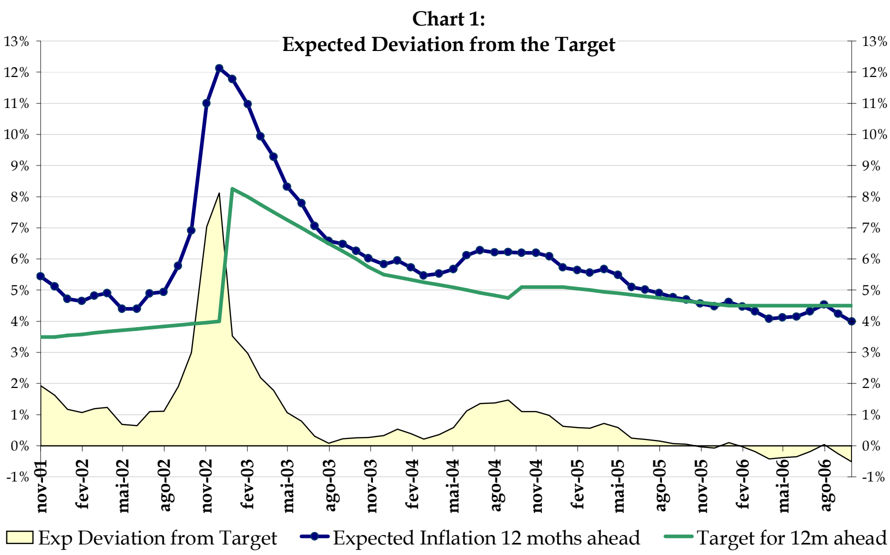
[4] There has been a history of changes in this policy objectives. When a big shock hits the economy and the target is no longer credible, the CB usually announce the objective that it will aim.

Year	Setting Date	Target		CPI Inflation	GDP Growth
1999	30/6/99	8.00%	± 2,00%	8.94%	0.79%
2000	30/6/99	6.00%	± 2,00%	5.97%	4.36%
2001	30/6/99	4.00%	± 2,00%	7.67%	1.31%
2002	28/6/00	3.50%	± 2,00%	12.53%	1.93%
2003	28/6/01	3.25%	± 2,00%	-	-
2003*	27/6/02	4.00%	± 2,50%	-	-
2003*	21/1/03	8.50%	± 2,50%	9.30%	0.54%
2004	27/6/02	3.75%	± 2,50%	-	-
2004*	25/6/03	5.50%	± 2,50%	7.60%	4.94%
2005	25/6/03	4.50%	± 2,50%	-	-
2005**	23/9/04	5.10%	-	5.69%	2.30%
2006	30/6/04	4.50%	± 2,00%	3.07%+	2.98%+
2007	22/6/05	4.50%	± 2,00%	4.15%'	3.49%'

* - Revised Targets.
** - Objective.
\+ - Consensus forecasts (means) on November 17, 2006

Figure 1:

months (solid green line), the inflation expectation for the next 12 months (dark blue line with circles) and the expected deviation from the target (filled area below), which is the difference between them:



But have short run inflation surprises been affecting these deviations from the announced target in Brazil? We proceed in an empirical investigation to estimate this effect. Define "inflation surprise" as the actual monthly inflation minus the expected inflation for this same month one month ago:

- "Inflation Surprise" $= CPI_t - E_{t-1}(CPI_t)$

In the Appendix we discuss this IPCA inflation surprise in more detail and show its daily-data dynamics. Also in the Appendix we present the a chart depicting the observed 12 month IPCA inflation, the target and the "12 month surprise".

Returning to our point here, the empirical evidence of co-movement between the 1 month inflation surprise and expected deviation from target 12 month ahead in Brazil is presented on Charts 2 and 3 below.

Chart 2:
Expected Deviation from Target vs. Inflation Surprise

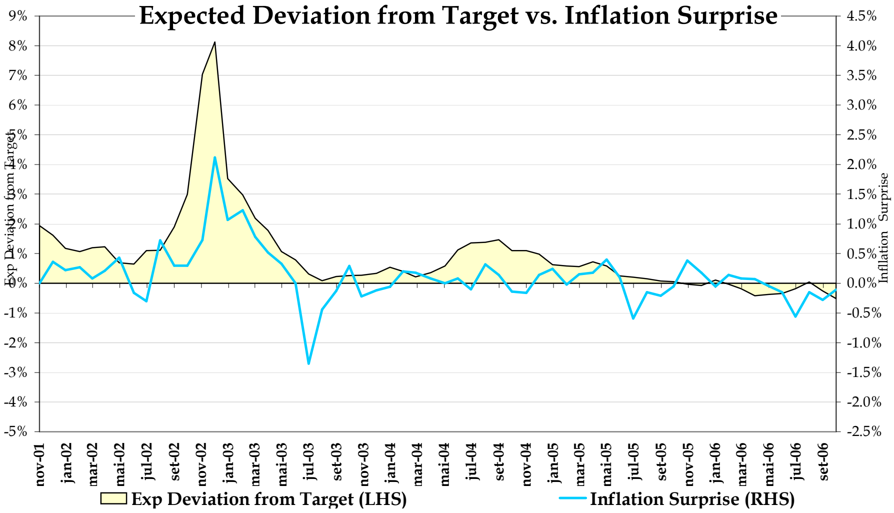
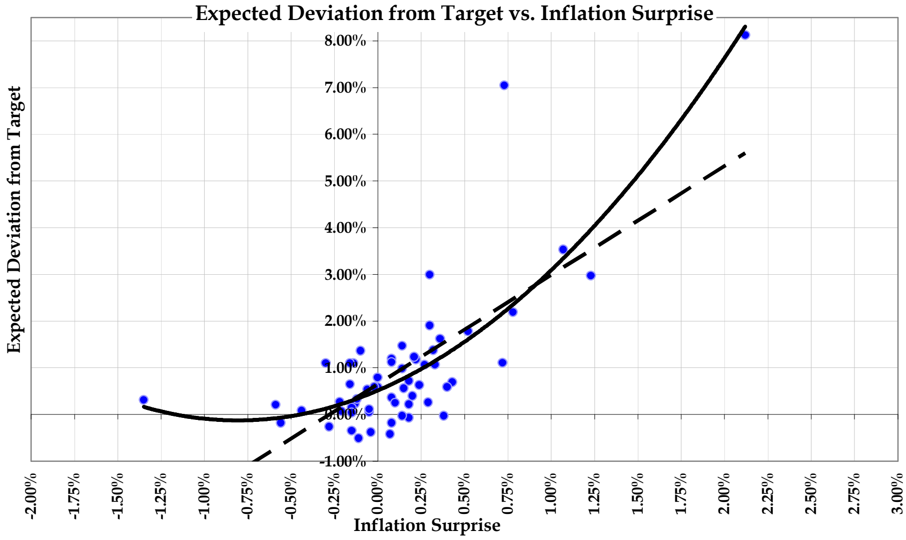


Chart 3:
Expected Deviation from Target vs. Inflation Surprise



What should be the effect of inflation surprise on the deviation from the target? Under perfect credibility and full commitment to the target, no effect should be observed. The story is different, however, the Central Bank (correctly) accommodates to some extent supply shocks. But even then, we should have a coefficient smaller than one since not all surprises are due to supply shocks. It is important to highlight that the measure of 12 month ahead inflation expectation that we use does not include the current month inflation. For this reason, under perfect credibility and no inflation inertia, the effect on inflation surprise on deviations from the target should always be zero even if the (current) supply shock is accomodated.

The figures above suggest that short run inflation surprises induce significant variation on medium run inflation expectation in Brazil. Linear and polynomial trends were included in the scatter plot and the polynomial one seem to provide a better fit. This indicates that the effect of inflation surprise on the expected deviation from the inflation target is non-linear: it increases with the size of the surprise and negative surprise seem to have negligible effect. This evidence goes against the null of good credibility of monetary policy or no inflation inertia in Brazil.

More careful econometric analysis is certainly needed to investigate that. To control for economic supply side pressures on inflation expectation, we included in our regressions the monthly exchange rate depreciation and also variation in the CRB commodity index. To control for demand side forces expected inflation, we use the CNI's seasonally adjusted installed capacity utilization. Since in the time period analyzed we had a severe confidence crises due to the 2002 elections, we look not only for the whole sample, but we also break it in different sub-periods even with our limited sample size[5].

The regressions presented below reinforce the impression we had by looking at the charts. When we run the regression for the whole period, the effect of inflation surprise on the expected IPCA deviation from the target is statistically significant, even controlling for exchange rate and commodity price effects with a coefficient of 0,77. An interesting fact is the asymmetric effect of positive and negative shocks in the whole sample: a 1% unexpected inflation shock leads to an increase of 1.75% on the 12 month inflation expectation above the CB's target while a negative surprise has no effect[6].

Results from regressions on the different sub-periods[7] indicates that the effect of short run inflation surprise on inflation expectation is diminishing over time. From Nov/2001 to Jun/2003 a 1% unexpected inflation shock provoked a 2,4% increase in expected deviation from the target. On the second subperiod, from Jul/03 to Dec/04, the effect diminished significantly as the effect of the same shock would lead only to an increase of 0,28% on 12 month ahead expected deviation from inflation target. On the more recent period, the effect is not significantly different from zero. Also interesting is the fact that the exchange rate effect on inflation expectations is also diminishing over time. The effects of a 10% depreciation on the expected inflation was around 1% on the first period, decreased to 0,5% in the second period and is now equals to zero.

[5] The sample is limited by the availability of data for inflation expectations, which started to be collected by the Brazilian Central Bank in 2001.

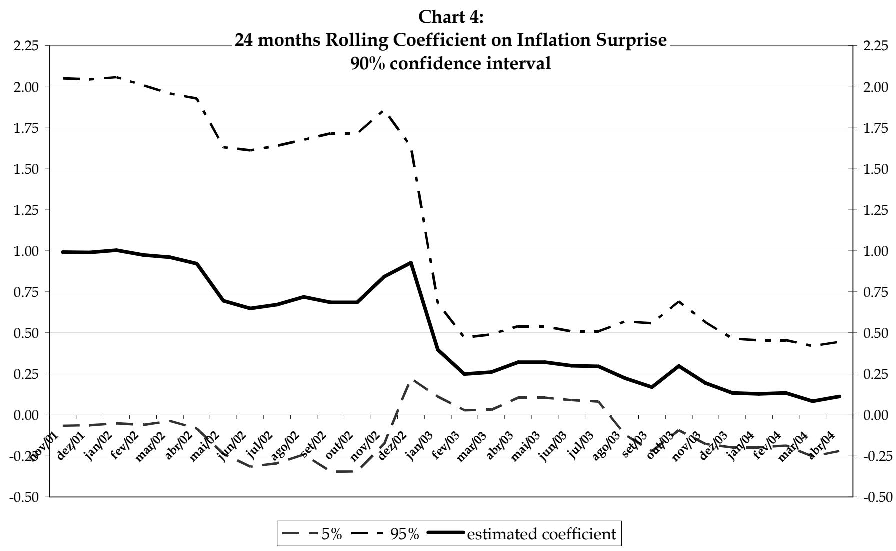
[6] This result is not robust to all sub-samples.

[7] In the sub-period regresions, we do include the control variables that didn't show up significant in the whole sample analysis because of the limited sample size.

OLS estimates	Dep Variable: Expected Deviation from Target = E(IPCA 12 month) - CB announced target 12 meses ahead							
	Eq. 1A	Eq. 1B	Eq. 2A	Eq. 2B	Eq. 3A	Eq. 3B	Eq. 4A	Eq. 4B
C	-0.010	-0.048	0.005	0.004	0.000	0.001	0.000	-0.001
	(0.879)	*(0.477)*	*(0.307)*	*(0.387)*	*(0.530)*	*(0.278)*	*(0.237)*	*(0.211)*
AR_1	0.665	0.488	0.233	0.225	0.996	0.998	0.909	0.898
	(0.000)	*(0.000)*	*(0.365)*	*(0.409)*	*0.000*	*0.000*	*0.000*	*0.000*
Surprise	**0.770**	-	**2.427**	-	**0.284**	-	-0.103	-
	(0.023)	-	***(0.019)***	-	***(0.022)***	-	*(0.510)*	-
Positive Surprise	-	**1.753**	-	**2.489**	-	-0.186	-	0.116
	-	***(0.002)***	-	***(0.032)***	-	*(0.677)*	-	*(0.762)*
Negative Surprise	-	-0.154	-	1.790	-	**0.389**	-	-0.229
	-	*(0.757)*	-	*(0.690)*	-	***(0.018)***	-	*(0.344)*
d(%XR(-1))	**0.068**	**0.072**	**0.098**	**0.095**	**0.051**	**0.044**	0.010	0.013
	(0.004)	***(0.002)***	***(0.039)***	***(0.066)***	***(0.008)***	***(0.027)***	*(0.303)*	*(0.237)*
d(%Commodities(-1))	0.011	0.017	-	-	-	-	-	-
	(0.771)	*(0.643)*	-	-	-	-	-	-
Capacity Utilization(-1)	0.000	0.001	-	-	-	-	-	-
	(0.857)	*(0.470)*	-	-	-	-	-	-
Adjusted R2	0.737	0.758	0.658	0.634	0.872	0.874	0.810	0.805
Durbin-Watson stat	1.704	1.479	1.465	1.412	1.857	1.811	1.521	1.620
AIC	-6.772	-6.840	-5.812	-5.708	-9.736	-9.716	-9.883	-9.823
Sample	Nov/01 - Oct/06	Nov/01 - Oct/06	Nov/01 - Jun/03	Nov/01 - Jun/03	Jul/03 - Dez/04	Jul/03 - Dez/04	Jan/05 - Out/06	Jan/05 - Out/06
N. Observations	*58*	*58*	*19*	*19*	*18*	*18*	*22*	*22*

P-values reported between parentheses

The choice of those 3 sub-periods was made in order to account for: (i) the pre and post electoral crises (Nov/01 - Jun/03), (ii) the re-built of the credibility under the new government (Jul/03 - Dec/04) and (iii) the present period (Jan/05 - Oct/06). Nonetheless, the choice of the specific dates for those periods was arbitrary. Hence, to better depict how this effect has been evolving over time, we also estimated 24 month rolling regression. The chart below present the results of the 90% confidence interval of the inflation surprise coefficient. Clearly, as previously noted, the effect of inflation surprise on deviation of inflation expectation over the target has diminished over time. Moreover, there seems to be 3 stable plateau levels for the coefficient: very high, lower but still significant and then low and insignificant.



To summarize, the overall evidence is that medium run inflation expectation was very sensitive to short run surprises in Brazil. More recently, apparently, this has not been a problem anymore. This does not mean, however, that it will not become again a problem in the future, as the Central Bank is not independent in Brazil yet.

For sake of comparison, we implement similar analysis with international data. Since we do not have the same disaggregation level available in international data[8] as we have in Brazil, the analysis is slightly different. Instead of using short run inflation surprises directly, we assume that the 1 month ahead inflation expectation is current month inflation, i.e., agents make projections as if monthly inflation were a random walk. We also modify the dependent variable, instead of using the expected deviation from the target we simply use the 12 month ahead inflation expectation, since some of the countries didn't have an announced target on our sample. We used market expectations survey and their sources are each country's central bank. We also run an instrumental variable regression where we use the lagged first difference of inflation as an instrument to the inflation surprise. The results are shown below:

OLS estimates	**Dep Variable: E(CPI 12 month/t) - E(CPI 12 month/t-1)**					
	CHILE	BRAZIL	TURKEY	UK	MÉXICO	ISRAEL
C	0.000	0.037	-0.903	-0.004	-0.086	-0.164
	(0.992)	*(0.824)*	*(0.012)*	*(0.840)*	*(0.052)*	*(0.322)*
d(Monthly Inflation(-1))	0.063	0.521	-0.044	-0.002	-0.052	0.143
	(0.325)	*(0.022)*	*(0.866)*	*(0.935)*	*(0.420)*	*(0.492)*
d(%XR(-1))	2.141	3.428	8.830	-1.200	2.488	15.979
	(0.027)	*(0.108)*	*(0.059)*	*(0.128)*	*(0.158)*	*(0.029)*
d(%Commodities(-1))	-1.611	-1.760	-13.070	1.554	1.073	-3.590
	(0.020)	*(0.628)*	*(0.161)*	*(0.112)*	*(0.329)*	*(0.734)*
R2	0.256	0.189	0.166	0.101	0.105	0.083
Durbin-Watson stat	1.438	1.146	1.394	2.075	1.832	1.111
AIC	-0.662	2.252	3.697	-0.887	-0.443	2.739
Sample	Oct/2001 - Oct/2004	Jan/2002 - Mar/2005	Oct/2001 - Nov/2004	Oct/1997 - Dec/2003	Jun/2001 - Sep/2004	Feb/1992 - Jan/1996
N. Observations	*37*	*39*	*38*	*75*	*40*	*48*

P-values in parenthesis

OLS estimates	**Dep Variable: E(CPI 12 month/t) - E(CPI 12 month/t-1)**					
	CHILE	BRAZIL	TURKEY	UK	MÉXICO	ISRAEL
C	0.000	0.024	-0.858	-0.006	-0.086	-0.139
	(0.990)	*(0.885)*	*(0.011)*	*(0.729)*	*(0.054)*	*(0.397)*
d(Monthly Inflation)	0.028	0.501	0.333	0.027	-0.016	-0.359
	(0.479)	*(0.062)*	*(0.084)*	*(0.416)*	*(0.846)*	*(0.061)*
d(%XR(-1))	2.583	2.701	8.221	-1.416	2.308	15.905
	(0.015)	*(0.136)*	*(0.072)*	*(0.178)*	*(0.200)*	*(0.012)*
d(%Commodities(-1))	-1.494	-0.790	-13.870	0.650	1.162	-7.472
	(0.041)	*(0.792)*	*(0.226)*	*(0.177)*	*(0.269)*	*(0.503)*
R2	0.219	0.178	0.219	0.064	0.096	0.123
Durbin-Watson stat	1.342	1.324	1.342	2.029	1.870	1.113
AIC	3.632	2.266	3.632	-0.846	-0.432	2.694
Sample	Oct/2001 - Oct/2004	Jan/2002 - Mar/2005	Oct/2001 - Nov/2004	Oct/1997 - Dec/2003	Jun/2001 - Sep/2004	Feb/1992 - Jan/1996
N. Observations	*37*	*39*	*38*	*75*	*40*	*48*

P-values in parenthesis

[8] We don't have 1 month ahead inflation to any country besides Brazil. Moreover, usually the data is for inflation expectaion for a given year, not for the following 12 months.

The results shown on the tables above suggest that in Brazil and Turkey[9] this effect is positive while in Chile, UK, Mexico and Israel there seems to be no effect at all.

Hence, the evidence so far is that short run inflation surprises induced a significant variation on medium run inflation expectation in Brazil (for the largest part of our sample), differently from most other countries. We conjectured that this phenomenon was happening for two (non mutually exclusive) reasons:

- Indexation of the economy.
- Lack of credibility of the Central Bank.

It is hard to argue that there is no remaining indexation in Brazil. A large part of the Brazilian CPI basket is composed of goods and services with managed prices. For instance, a quick look at public concession contracts such as telephony and energy services reveals that inflation indexation is quite alive: they are adjusted annually by the previous 12m inflation.

What we are interested in is in figuring out if, besides inflation inertia, the lack of credibility was also present. If this is the case, the benefit of an independent Central Bank would be clear. For the endeavor, what we propose in the next section is a methodology to identify if the phenomenon itself is somehow related to the lack of credibility of the Central Bank and then, in the fourth section, we apply it to the Brazilian data. The idea is to track *"Inflation Risk Premium"* movements.

3 The Model

Let us start by formally defining the "break even inflation" (also called "market inflation") and the "inflation risk premium":

- "Break Even Inflation" = (1+Nominal Rate)/(1+ Real Rate)-1
- Inflation Risk Premium = (Break Even Inflation) - (Expected inflation)

The main message of the model is the following: If the cause of the effect of short run inflation surprise on 12 month inflation expectation is solely indexation, there is no reason for an increase in the uncertainty when the economy is hit by a positive inflation shock: we know that the prices will be re-adjusted in the future with certainty. However, if there is lack of credibility on monetary policy, there will be an increase in the uncertainty on future responses to inflation, leading to an increase in the uncertainty on inflation itself. This will be captured by the inflation risk premium. We now formalize this argument.

[9] In Turkey the effect is not observed when we run the instrumental variable regression.

3.1 Asset Pricing and the Inflation Risk Premium

Define an economy with two assets: a nominal bond ($P_t^{\$}$) and a real bond ($P_t$). These bonds will be freely traded in the market. In order to price them, we will impose absence of arbitrage. This imply that there is a strictly positive stochastic discount factor M_{t+1} such that for any stochastic *nominal payoff* X_{t+1} to be observed in $t+1$, its time t price will be given by $P_t = E_t\left[M_{t+1}X_{t+1}\right]$. If markets were complete, the stochastic discount factor used to price any asset in this economy would be the same. We don't need that much, we can have incomplete markets since the M_{t+1} used to price the real bond can also be one of the stochastic discount factors used to price the nominal bond.

In this setting, if the *nominal payoff of the real bond is* Π_{t+1}, i.e., (1+inflation%), and the *nominal payoff of the nominal bond is* \$1, their prices will be:

- Real Bond: $P_t = E_t\left[M_{t+1}\Pi_{t+1}\right]$
- Nominal Bond: $P_t^{\$} = E_t\left[M_{t+1}\$1\right]$

If an investor buys a nominal bond he will not be hedged against inflation, so a nominal bond needs to compensate investors for inflation risk. Notice that if $E_t\Pi_{t+1} > 1$, then $P_t^{\$} < P_t$ or, in other words, the real rate will be smaller than the nominal rate.

The log real rate is $r_t \equiv -\ln P_t$ and the log-nominal rate is $r_t^{\$} \equiv -\ln P_t^{\$}$. We can specify the process of the short rate and the prices of risk which is equivalent to specifying a process for the stochastic discount factor. Our main hypothesis is that this short rate process is defined by the Central Bank and can be characterized by a state-dependent "Taylor rule". This is already usual in the recent Macro-Finance term structure literature and is also done in Ang & Piazzesi (2003), Rudebusch & Wu (2004) and Bonomo & Lowenkron (2006). Define the Taylor rule as:

$$r_t = \delta^0 + \delta_t^1 X_t \tag{1}$$

In (1) X_t are the state variables of the economy that affects the decision of the Central Bank, such as current inflation and output. Notice that we allow the Central Bank response to state variable shocks to be time-varying[10]. Let the dynamic of the state variables be driven by a VAR process:

$$X_t = \mu + \Phi X_{t-1} + \varepsilon_t$$

We also suppose that those sources of uncertainty on the economy are associated with a vector λ containing the prices of risk of each state variable of the economy. We follow Vasicek (1977) and suppose

[10] Think of that in an monetary policy framework where the Central Bank is not independent: there is a possibility that, for electoral reasons, the cuntry's president can change the staff of the Central Bank, generating a change in its Taylor rule.

that the prices of risk are not time varying[11]. To use the pricing formulas above, we must find the stochastic discount factor and the connection between the short rate and the stochastic discount factor is given by the Radon-Nikodym[12] derivative ξ_{t+1} that changes the probability measure to a martingale equivalent one:

$$\xi_{t+1} = \xi_t \exp\left(-\frac{1}{2}\lambda'\lambda - \lambda'\varepsilon_{t+1}\right) \tag{2}$$

Recall that ε_{t+1} is the vector of uncertainties affecting the state variables X_{t+1}.

The stochastic discount factor (or pricing kernel) will be given by:

$$M_{t+1} = \exp\left(-r_t\right)\frac{\xi_{t+1}}{\xi_t} \tag{3}$$

Substituting eq. (1) and eq. (2) in eq. (3), we get:

$$m_{t+1} \equiv \ln M_{t+1} = -\frac{1}{2}\lambda'\lambda - \delta^0 - \delta_t^{1'} X_t - \lambda'\varepsilon_{t+1} \tag{4}$$

Now we can calculate the prices of the relevant securities. Our task is to understand what affect the inflation risk premium, which is defined by the difference between the log "Break Even Inflation" ($\log \Pi_{t+1}^{BE}$), and the agent's log expected inflation ($\log E_t(\Pi_{t+1})$). The "break even inflation" is simply the inflation rate implicit on the financial securities, and it can be measured by the nominal forward rate minus the forward real rate $\log \Pi_{t+1}^{BE} = f_{t+1}^{\$} - f_{t+1}$. So, these are the securities that we are interested in to calculate the inflation risk premium.

Define $m_{t+1} \equiv \ln M_{t+1}$, $\pi_{t+1} \equiv \ln \Pi_{t+1}$ and suppose that M_t and Π_{t+1} are jointly lognormally distributed. In this setting, the price of the 1 period nominal bond will be given by:

$$p_t^{1\$} = E_t\left(m_{t+1}\right) + \frac{1}{2}Var_t\left(m_{t+1}\right) \tag{5}$$

And the price of the real bond will be given by:

$$p_t^1 = E_t\left(m_{t+1}\right) + E_t(\pi_{t+1}) + \frac{1}{2}Var_t\left(m_{t+1}\right) + \frac{1}{2}Var_t(\pi_{t+1}) + Cov_t(m_{t+1}, \pi_{t+1}) \tag{6}$$

In the same way, we can also calculate the price of the 2 period nominal and real bonds[13]. Using these formulas, we calculate the 1 period real forward rate, which is given by $f_t = p_t^1 - p_t^2$ and the 1 period nominal forward rate, given by $f_t^{\$} = p_t^{1,\$} - p_t^{2,\$}$:

$$f_t^{\$} = -E_t\left(m_{t+2}\right) - \frac{1}{2}Var_t\left(m_{t+2}\right) - Cov(m_{t+1}, m_{t+2}) \tag{7}$$

[11] Later, we will discuss what would be the implications of relaxing this hypothesis. We will argue that, for our purpuses, it wouldn't make much difference.

[12] Radon-Nikodym derivative changes the probability measure to a maringale equivalent one. However, to price assets, it is equivalent to work with the present value of the expected payoff under the martingale equivalent measure or the expected payoff multiplied by the stochastic discount factor.

[13] This calculation is shown in the appendix.

$$
\begin{aligned}
f_t \quad = \quad & -E_t\left(m_{t+2}\right) - E_t(\pi_{t+2}) - \frac{1}{2}Var_t\left(m_{t+2}\right) - \frac{1}{2}Var_t(\pi_{t+2}) - \\
& -Cov(m_{t+1}, m_{t+2}) - Cov(\pi_{t+1}, \pi_{t+2}) - \\
& -Cov_t(m_{t+1}, \pi_{t+2}) - Cov_t(m_{t+2}, \pi_{t+1}) - Cov_t(m_{t+2}, \pi_{t+2}) \qquad (8)
\end{aligned}
$$

To calculate the inflation risk premium we also need the expected log inflation that under the log-normality assumption is given by:

$$\log E_t(\Pi_{t+1}) = E_t(\pi_{t+1}) + \frac{1}{2}Var_t(\pi_{t+1})$$

The formula of the 2 period inflation $\Pi_{t+2} = e^{\pi_{t+1}+\pi_{t+2}}$ is given by:

$$\log E_t(\Pi_{t+2}) = E_t(\pi_{t+2}) + E_t(\pi_{t+1}) + \frac{1}{2}Var_t(\pi_{t+1}) + \frac{1}{2}Var_t(\pi_{t+2}) + Cov(\pi_{t+1}, \pi_{t+2})$$

In turn, the expectation of the log 1 period inflation in $t+2$ will be:

$$\log E_t(\Pi_{t+2}/\Pi_{t+1}) = E_t(\pi_{t+2}) + \frac{1}{2}Var_t(\pi_{t+2}) + Cov(\pi_{t+1}, \pi_{t+2}) \qquad (9)$$

Now we are ready to calculate the inflation risk premium substituting out the equations (7), (8) and (9) on the following definition:

$$
\begin{aligned}
\textit{Inflation Risk Premium}_t \quad \equiv \quad & \log \Pi_{t+1}^{BE} - \log E_t(\Pi_{t+2}/\Pi_{t+1}) \\
= \quad & f_t^{\$} - f_t - E_t(\pi_{t+2}^f) \\
= \quad & Cov_t(m_{t+1}, \pi_{t+2}) + Cov_t(m_{t+2}, \pi_{t+1}) + Cov_t(m_{t+2}, \pi_{t+2}) \qquad (10)
\end{aligned}
$$

To relate the Taylor rule with inflation risk premium, we still need to substitute the s.d.f. equation (4). We also get rid of constant terms, that are irrelevant for the covariance, and arrive at:

$$\textit{Inflation Risk Premium}_t \equiv Cov_t(-\delta_t^{1'} X_t - \lambda'\varepsilon_{t+1}, \pi_{t+2}) + Cov_t(-\delta_t^{1'} X_{t+1} - \lambda'\varepsilon_{t+2}, \pi_{t+1}) + Cov_t(-\delta_1' X_{t+1} - \lambda'\varepsilon_{t+2}, \pi_{t+2}) \qquad (11)$$

In the next section we will impose some further restrictions to get a more tractable formula for the inflation risk premium.

3.2 Taylor Rule, Inflation Dynamics and the Inflation Risk Premium

Let us recall the main motivation of the paper. The empirical evidence presented in Section 2 was that short run inflation surprises induced a significant variation on medium run inflation expectation in Brazil, differently from most other countries. We conjecture that this phenomenon could be happening for two (non mutually exclusive) reasons:

- Indexation of the economy.

- Lack of credibility of the Central Bank.

As said before, it would be hard to argue that there is no remaining indexation in Brazil. What we are interested in is in figuring out if the lack of credibility is also present. The strategy that we are suggesting is to look at the effects of inflation surprises in inflation risk premium, which brings information about future responses to inflation shocks. In order to have more clear equation determining inflation risk premium, we will further parametrize and impose more structure in the model.

We start with the equation for inflation dynamic. A natural question that arises is: Can inflation inertia be the responsible for variations in inflation risk premium? For this reason, in the model we will allow inflation to have an inertia component. Assume that the log inflation has a certain degree of persistence, a long run average and that it is affected by monetary policy. The equation proposed is given by:

$$\pi_t = (1-\phi_\pi)\mu_\pi + \phi_\pi \pi_{t-1} - \phi_r(r_{t-k} - \overline{r}) + \varepsilon_t^\pi \tag{12}$$

Where, μ_π is the long-run "natural inflation"; ϕ_π is the inflation inertia (degree of indexation of the economy); ϕ_r is the inflation sensitivity to monetary policy, $\varepsilon_t^\pi \sim N(0,\sigma)$ is the inflation shock at time t and; k is the lag with which the monetary policy affects the economy.

Now we turn to the Taylor rule. We permit for central bank's response to be time varying. Think of that as the possibility of changes in the board of the Central Bank. Without Central Bank independence, the government can appoint the CB board according to its will.

Since we already know that all that matters for inflation risk premium is the covariance of inflation with the stochastic discount factor, we just need to worry about the Central Bank responses to inflation shocks without any loss generality. Recall that ε_t^π is the unexpected inflation shock at time t. We will allow the central bank to respond to contemporaneous inflation shocks (ε_t^π) and/or any past inflation shocks ($\varepsilon_{t-j}^\pi,_{for\ j\geq 1}$), with different elasticities of response to each one of them. The idea is that with this formulation, we could account for as many kinds of monetary rules as possible. The state-dependent Taylor Rule will be given by:

$$r_t = \overline{r} + \delta_0^\theta \varepsilon_t^\pi + \delta_1^\theta \varepsilon_{t-1}^\pi + \delta_2^\theta \varepsilon_{t-2}^\pi + ... \tag{13}$$

Where $\overline{r}$ is the long run "natural" interest rate. δ_t^θ is the type θ monetary authority policy response[14] to inflation shock at time t.

[14] Should the monetary authority react to current inflation shocks ε_t^π raising interest rates if it wants to diminish inflation, i.e., should $\delta > 0$? If the $k = 0$ the answer is clearly yes. But it is also easy to see that even if the monetary policy affect the economy with some lag ($k > 0$), the central bank will want to have $\delta > 0$ if we have some inflation inertia ($\phi_\pi > 0$).

Now, we are ready to find a more simple expression for inflation risk premium expression (11) substituting in it the Taylor rule equation (13) and the inflation dynamics equation (12):

$$
\begin{aligned}
\textit{Inflation Risk Premium}_t &= \left[-\lambda - \delta_o\right](1+\phi_r)\sigma^{\pi} \quad \text{if k=0} \\
&= (-\lambda - \delta_o)\sigma^{\pi} \quad \text{if k>0}
\end{aligned} \tag{14}
$$

This is enough to see that indexed economy does not provoke a positive relation between inflation surprises and inflation risk premium, since the indexation parameter ϕ_π does not appear in equation (14) neither in the more general equation (11)[15].

Would could cause variations in the inflation risk premium then? There are only 3 possible causes[16] of variations in inflation risk premium: (1) variations in the conditional expected volatility of future inflation σ^{π}_{t+1}; (2) variations in the price of risk of inflation λ; (3) variations in the expected central bank response to future inflation shocks δ^1_{t+1}. Notice that here we have even relaxed the hypothesis that market price of risks are constant. Therefore, these are the only three possibilities without imposing any further assumption[17].

On this paper, we will focus only on the third possible explanation: variations in the expected CB responses to inflation. The basic insight is already in equation (14): consider two types of Central Bank, a tougher one that responds aggressively to a contemporaneous inflation shock (high δ^1) and a second type, that has a loose monetary policy and do not respond so much to contemporaneous inflation shock (low δ^1). The tight monetary authority type will be associated with a lower inflation risk premium.

Suppose that there can be two types of Central Bank: (i) one more committed to fight inflation, i.e, with a higher interest rate response to inflation shock δ^H; and (ii) one less committed to fight inflation, i.e, with a lower interest rate response to inflation shock δ^L. Moreover, let p_t denote the conditional probability at time t of the type of monetary authority in $t+1$ being L and $(1-p_t)$ the the conditional probability at time t of the type of monetary authority in $t+1$ being H.

Now we can have a clearer expression to equations (10) and (11):

$$
\begin{aligned}
\textit{Inflation Risk Premium}_t &= \left[p_t(-\lambda - \delta^L_o) + (1-p_t)(-\lambda - \delta^H_o)\right](1+\phi_r)\sigma^{\pi} \quad \text{if k=0} \\
&= \left[p_t(-\lambda - \delta^L_o) + (1-p_t)(-\lambda - \delta^H_o)\right]\sigma^{\pi} \quad \text{if k>0}
\end{aligned} \tag{15}
$$

[15] If the indexation was time-varying then this term would be present in equation (14), thus this result is not robust to the that. However, we think that indexation of an economy is a very low frequency characteristc of an economy and, therefore, should not affect the result.

[16] Actually, if one go back to equation (11) will see that the relevant covariance is the one the state variables in the vector X_t. Inflation π_t is certainly one of them, but if there were contemporaneous covariance between π_t and other state variables to which the Central Bank react to through its Taylor rule, there would be other possibilities besides those 3.

[17] The only necessary hypothesis are: (i) absence of arbitrage, (ii) linear Taylor rule and (iii) linear price of risk and (iv) the contemporaneous covariance of inflation with other variables entering the Taylor rule is negligible.

According to our formulation, if we observe a positive relation between inflation surprises and inflation risk premium, what is happening is that a positive inflation shocks is inducing an increase in the perceived probability p_t that that the monetary policy next period will be more loose than the current one.

What could induce that perception? Recall that we supposed that we do not have Central Bank independence, so the chief of government (president) can choose to change the board according to his will. As an election approaches, if the economy is hit by negative shocks (such as negative shocks) there will be an increase in the probability in the change in the governing party. If moreover the new party have different preference concerning how hard should on fight inflation by increasing interest rates, a variation in risk premium will be observed.

Another possibility is strategic behavior by the incumbent government. Let us suppose that there is a short run trade off between inflation and output gap as in a Phillips curve. Suppose also that output gap affects voters evaluation of the incumbent government negatively and politicians care only about being elected as in an "opportunistic"[18] model in the spirit of Persson and Tabellini (1990) and Rogoff and Silbert (1988). These models have a non-observable competence term. The voters don't know if the better outcome was achieved by exploiting the Phillips curve or by a higher competence of the government so they vote for the incumbent when output is good (or at least not too bad). A positive relation between unexpected inflation shocks and inflation risk premium reflects financial agents expecting a change in the conduction of monetary policy towards a not so tight one, since the scenario is not so benign to the incumbent government. In turn, this happens because due to electoral concerns, the president can change the board of the Central Bank if the state of the nature (inflation shocks) is not in his favour. If Central Bank were independent, certainly this problem wouldn't happen.

To reconcile that last reasoning with our result that some agents are predicting a change in future responses to inflation shocks (change in the Taylor rule) we would need to introduce asymmetric information. Suppose that the agent who trades in the financial market and determine prices is not the *median voter*. This last hypothesis is done in Mankiw and Zeldes (1991): they observe that only a fraction of the families invest on the financial market in the USA and then use that to try to solve the equity premium puzzle[19]. We claim that it makes sense (and is also legitimized in the literature) to use different representative agents: one to price the assets and the other to elect the government. So think as our result under this political economy reasoning as if positive relation between unexpected inflation shocks and inflation risk premium reflects financial agents expectation, not the median voter expectation.

[18] Nordhaus (1975) and Lindbecck (1976) introduced this approach but in their model, voters are naïve and the agents do not have rational expectations.

[19] The volatitlity of consumption of equity investors is much higher than the average US resident, so in a CCAPM model the s.d.f. is more volatile, requiring a much smaller risk aversion to calibrate for the data on equity returns.

To conclude, in our formulation, if inflation risk premium is changing in the presence of positive unexpected inflation shocks agents are fearing that the "type" of the monetary policy can change. Precisely, agents are fearing a loosening of monetary policy $\delta^L < \delta^H$ when the economy is hit by a positive inflation shock, causing an increase in inflation risk premium[20]. This fact has a deleterious effect on economy since the inflation expectations should also increase in the presence of this shock, what makes monetary more costly in terms of output loss. It should be clear that we are not saying that lack of credibility of the monetary policy is the only factor: inflation indexation can be important too[21]. Thus, if the the empirical pattern indicate this positive relation this is an indication that there would room to reduce the cost of monetary policy by promoting Central Bank independence.

4 Inflation Risk Premium in Brazil

What could be causing the positive relation between short run inflation surprise and medium run inflation expectation as identified in Brazil in section 2? Does monetary policy credibility plays a role in that? The results from last section indicates that looking at the relation between inflation surprise and inflation risk premium helps to determine if the phenomenon is somehow related to the lack of credibility of the Central Bank. In this section we investigate if this is present in the Brazilian data.

The intuition of the model was that if the cause of the effect of short run inflation surprise on 12 month inflation expectation is solely indexation, there would be no reason for an increase in uncertainty when the economy is hit by a positive inflation shock: we know that the prices will be readjusted in the future with certainty. However, if there is lack of credibility on monetary policy, there would be an increase in the uncertainty on future responses to inflation which, in turn, would lead to an increase in the uncertainty on inflation itself. This will be capture by the inflation risk premium.

Now we look at the empirical relation between inflation risk premia and inflation surprises in Brazil. There was no liquid market for bonds indexed to IPCA until recently, so we had to resort to a different inflation index: IGPM[22]. All over the time frame analyzed, there was a liquid market for government bonds indexed to this inflation indexed, the NTN-Cs. The data used in the analysis is the interpolated 1 year real rate and 1 year nominal rate[23], both provided Brazil's Future and Mercatile Exchange (BM&F) and the survey of market expectation for IGPM 1 year inflation, which was obtained again in Brazilian

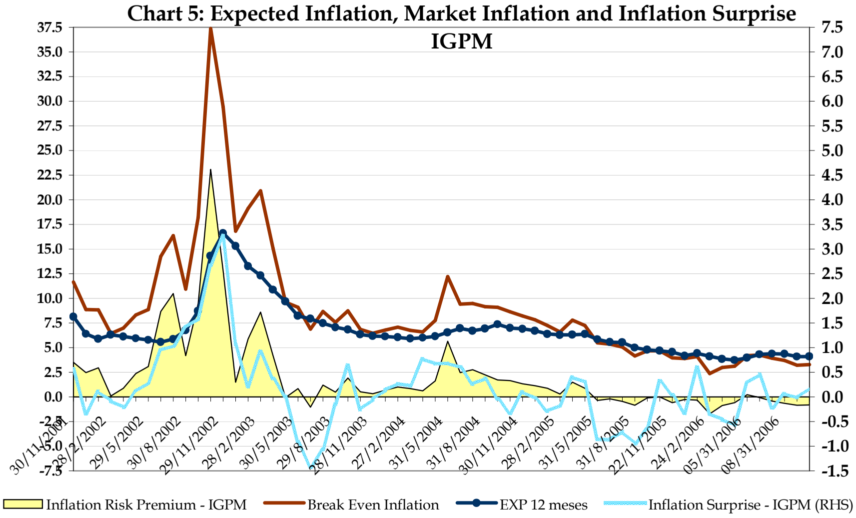
[20] An important observation made by Ilan Golfaijn is that the expected loosening of the monetary policy do not necessarily have to be associated with a lack of credibility of the monetary authority. If agents expect that during some period, supply shocks will be the dominant type of shocks, they can rationally expect a smaller interest rate response than before. And this response by the monetary authority is compatible with the optimal response, as in Woodford (2002).

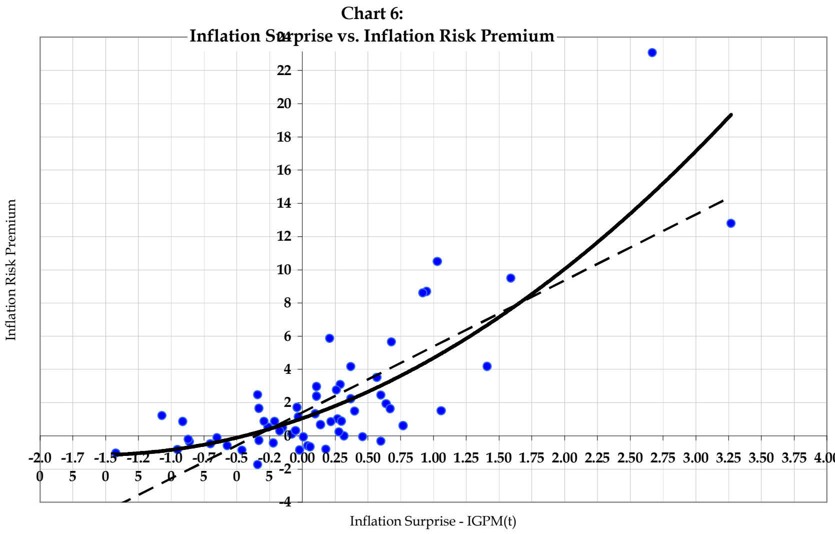
[21] And as said before, we think it is as the examples of the public services contracts such as telephones and energy points out.

[22] This is a general price index, basically an average of producer price index and consumer price index.

[23] Swap DI-pré 1 year.

Central Bank survey database. Below, we present graphs of the evolutions of these series. The brown line is the break even inflation, the dark blue line with circles is the 12 month expected inflation. The difference between these two is the Inflation risk premium, represented by the shaded area. The short run inflation IGPM surprise[24] is the light blue line. We also present scatter plot:





As in the case of expected deviation from the target analyzed in section 2, we also fin here a clear pattern of positive relation between inflation surprise and inflation risk premium. As stressed by our model, this indicates that the episode documented in section 2 is more likely to be associated with

[24] Calculated just like the the IPCA surprise.

expectation of more dovish monetary policy. As before, we continue our investigation with a series of regressions, which are presented below:

OLS estimation	**Dep Variable: IGPM Inflation Risk Premium**							
	Eq. 5A	Eq. 5B	Eq. 6A	Eq. 6B	Eq. 7A	Eq. 7B	Eq. 8A	Eq. 8B
C	-0.794	-1.199	3.444	3.154	0.436	0.813	-1.475	-1.407
	(0.17)	*(0.03)*	*(0.21)*	*(0.18)*	*(0.81)*	*(0.73)*	*(0.01)*	*(0.01)*
AR_1	-0.008	-0.152	-0.532	-0.755	0.211	0.221	0.329	0.374
	(0.95)	*(0.24)*	*(0.09)*	*(0.01)*	*(0.38)*	*(0.38)*	*(0.09)*	*(0.08)*
Surprise	**2.276**	-	**5.908**	-	**1.092**	-	**0.631**	-
	(0.00)	-	***(0.00)***	-	***(0.08)***	-	***(0.01)***	-
Positive Surprise	-	**4.116**	-	**8.207**	-	1.478	-	0.936
	-	-	-	***(0.00)***	-	*(0.36)*	-	*(0.13)*
Negative Surprise	-	0.336	-	2.765	-	0.714	-	0.427
	-	*(0.69)*	-	*(0.17)*	-	*(0.65)*	-	*(0.33)*
d%XR	12.475	10.191	6.728	2.205	12.412	13.626	**5.895**	**6.501**
	(0.10)	*(0.15)*	*(0.70)*	*(0.88)*	*(0.33)*	*(0.33)*	***(0.03)***	***(0.03)***
d%Commodities	-5.147	-7.920	-38.431	-44.023	-3.432	-2.591	**5.952**	**5.820**
	(0.61)	*(0.40)*	*(0.19)*	*(0.09)*	*(0.78)*	*(0.85)*	***(0.05)***	***(0.06)***
Embi+ Brazil	**0.372**	**0.329**	0.135	0.089	0.119	0.024	**0.423**	**3.000**
	(0.00)	***(0.00)***	*(0.57)*	*(0.66)*	*(0.68)*	*(0.96)*	***(0.01)***	***(0.04)***
Adjusted R2	*0.743*	*0.779*	*0.725*	*0.800*	*0.219*	*0.159*	*0.734*	*0.722*
D-W	*1.995*	*1.924*	*2.093*	*2.406*	*2.262*	*2.256*	*1.783*	*1.786*
AIC Criteria	*4.398*	*4.262*	*5.334*	*5.042*	*3.441*	*3.540*	*1.033*	*1.103*
Sample	**Nov/01 - Oct/06**	**Nov/01 - Oct/06**	**Nov/01 - Jun/03**	**Nov/01 - Jun/03**	**Jul/03 - Dez/04**	**Jul/03 - Dez/04**	**Jan/05 - Out/06**	**Jan/05 - Out/06**
N. Observations	*58*	*58*	*19*	*19*	*19*	*19*	*22*	*22*

P-values between brackets

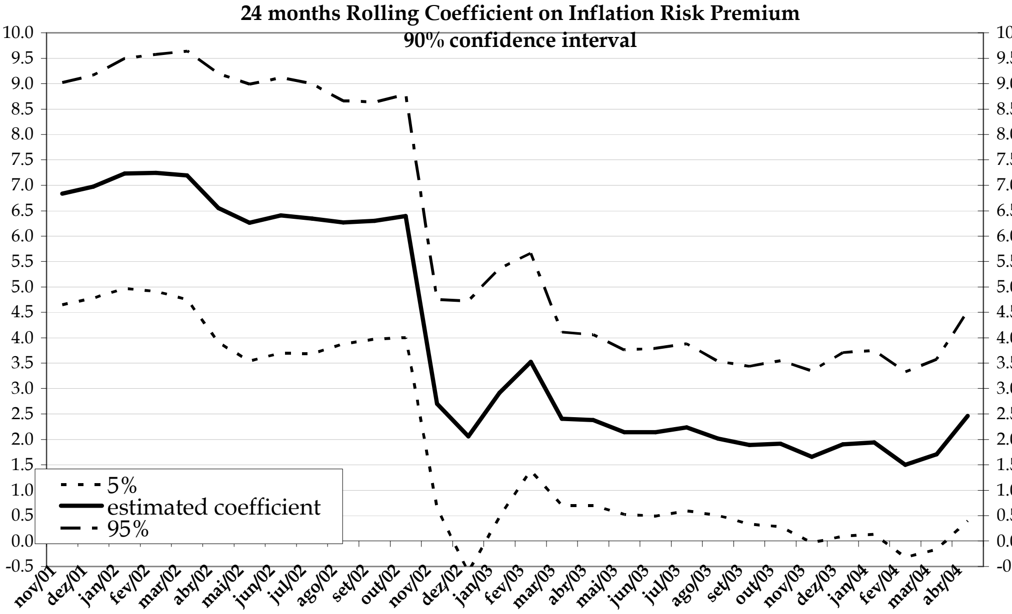
The coefficient on inflation surprises is positive and significant in all specifications. In all of the estimated equations the magnitude is much greater than the ones in the inflation expectation regressions. On average a 1% inflation surprise increase inflation risk premium in 2.28% (Eq. 5A). Asymmetric effects are also present in the whole sample analysis (Eq. 5B) as a 1% positive surprises have 4.11% effect on inflation risk premium and a negative surprise have an insignificant effect .

Interestingly, the EMBI spread also have a positive effect in inflation risk premium but its inclusion don't change the significance of inflation surprises[25]. We could interpret that as accounting for variations in the risk aversion, the parameter λ in our model.

As on the previous analysis, the effect has clearly been diminishing in recent time. This can be seen by the evolution of the surprise coefficient in regressions. The effect of 1% inflation surprise in the first sample period was 5.9% (eq. 6A), became 1.0% in the second sample period (eq. 7A) and is right now very small, but still significant, at a 0.6% level (eq. 8A). The choice of the sample periods had the same objective as in section 2 but as they were arbitrary, we also plot the result of 24 month rolling regression in the Chart 7 below:

[25] Recall that earlier we said that one of the possible causes of the time variation in inflation risk premium is a time varying price of risk. Embi is the usually the best instrument for risk aversion in Brazil and its inclusion does not change the significance of short run inflation surprises. Thus, we believe that the effect that we are focusing in, namely the expected change in Taylor rule, is indeed present.

Chart 7:
24 months Rolling Coefficient on Inflation Risk Premium
90% confidence interval



So not only the expected deviation of the inflation from the targeting was responding to short run inflation surprises as we already knew from section 2 results. The evidence presented in this section is that at the same time, inflation risk premium was responding very strongly as well to those shocks. According to the model proposed in section 3, the interpretation to this fact is that agents did not believe that the same interest rate response to inflation shocks would continue in the following 12 months: a relaxation in the fight against inflation was expected. Possible reasons for that were already discussed.

Variations in the expected monetary authority response to future inflation shocks, i.e., the credibility of the ongoing policy, is certainly not the only possible explanation for the phenomenon we have documented here. Our general model also suggested that another possibility for time-varying inflation risk premium is the presence of heteroskedasticity in unexpected inflation shocks. But economically what would this mean? We believe that this could also be (but not necessarily would be) associated with lack of credibility in the monetary authority as big surprises would be followed by big surprises. And those surprises could be big precisely because of a confidence crisis as we had in 2002 in Brazil. The third and last possibility, also according to our general model, would be a time varying market price of risk for inflation λ. In the regressions we allowed the inflation risk premium to depend on the best proxy we found to a "price of risk", the EMBI+ Brazil spread. If that is a good measure of price of risk, its inclusion did not drive out the importance of inflation surprises. Therefore we conclude that variations in the expected monetary policy response to inflation shocks was important in Brazil during the time period analyzed. This is an indication that imperfect credibility in the monetary arrangement contributed to to clogging a transmission channel of the monetary policy, therefore imposing a higher equilibrium real rate than otherwise.

5 Conclusion

Short run inflation surprises has been affecting medium run inflation expectations in Brazil. This is a symptom of at least one of two problems: (i) Inflation inertia / indexation of the economy; and/or (ii) lack of credibility of the monetary authority. The remedy depends on the cause. For instance, if the reason is simply indexation, central bank independence would not help.

We presented a model arguing that movements in inflation risk premium can help to identify if credibility is one of the causes of this phenomenon. Empirical results indicate that in Brazil, inflation surprises have pushed expected inflation away from the target and have also driven inflation risk premium up for most of our sample. This is an indication that imperfect credibility in the monetary arrangement was clogging an important transmission channel of the monetary policy and, therefore, imposing a higher equilibrium real rate than otherwise. We conclude, therefore, that had central bank independence in place during this period, it would have helped monetary policy in Brazil to be significantly less costly. Other evidence that points in this direction is that bad news (positive inflation shocks) had more effect than good news on both medium run inflation expectation and inflation risk premium in our sample. Recently all of these effects of inflation surprise on inflation expectation and risk premium have diminished and eventually disappeared, what we interpret as an indication of improvement in monetary policy credibility - possibly being one of the responsibles for the lower real rate we are observing today. This does not mean, however, that it will not become again a problem in the future, as Central Bank is not independent in Brazil yet.

We conjecture that the lack of credibility can not be understood looking at the history of conservative decisions of the Brazilian central bank in our sample. As we put in our model, what we believe that harms the credibility is the fear of regime switch (peso problem). In line with political economy literature[26], the reasoning for goes as follows: bad news on inflation would require an even tougher monetary policy, reducing the chance of reelection of the incumbent party. In such a scenario, constantly arriving good news would be necessary to prevent a regime switch and that's why bad news would have such a deleterious effect.

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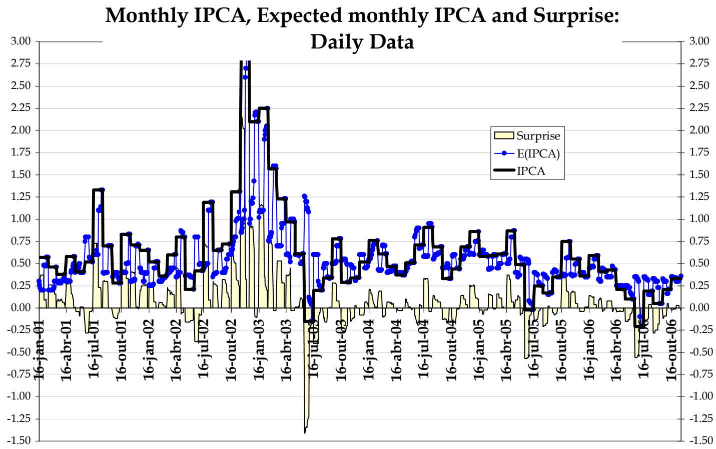
7 Appendix

7.1 Monthly IPCA surprise: daily data and methodology

The Chart below depicts the dynamic of the surprise on monthly IPCA. To better understand what is going on, it is useful to first discuss some methodological issues of the measures involved.



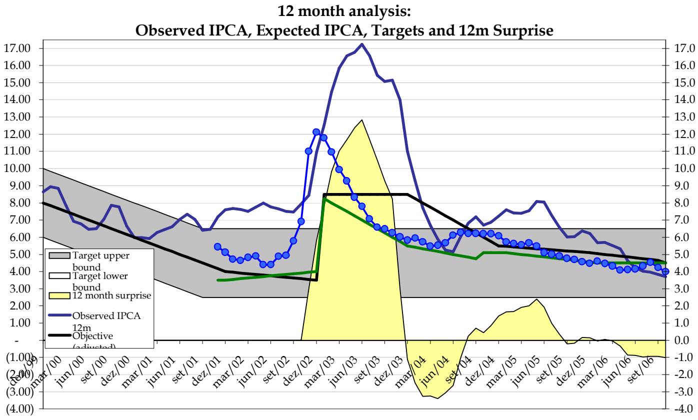
The month t IPCA is announced on the day 15th of month $t+1$ and refers to the inflation from the first to the last day of that previous month. On the first business day of each month, the same statistics office responsible for this inflation index, IBGE (Instituto Brasileiro de Geografia e Estatística), also report the IPCA-15 in which the basket of goods and services is the same one of the IPCA. The only difference between them is that IPCA-15 refers to the inflation from the day 15 form month t to day 15 from month $t+1$.

The black solid line in the chart represents the next monthly IPCA actually announced. For this reason, on the 16^{th} of each month, we observe a jump and it stay in the new level until the 16th of the following month. The blue line with circles are the market expectation for the next month IPCA. The surprise is the difference between the two measures and is represented by the filled area.

As one would expect, the measure of surprise diminishes we approach the announcement day. However, not surprisingly, the biggest change occurs around the 1st day of each month, when the IPCA-15 is announced. From the 16^{th} day to the 1^{st} day of each month, the change in the expected IPCA for that month is very smooth.

For those reasons and since we are working with monthly data, the measure of surprise we use in this paper is the point around the 16^{th} of each month.

7.2 Accumulated 12 month IPCA surprise



7.3 Pricing formula for 2 period bonds:

The price of a 2 period bonds:

$$p_t^{2\$} = E_t\left(m_{t+1}\right) + E_t\left(m_{t+2}\right) + \frac{1}{2}Var_t\left(m_{t+1}\right) + \frac{1}{2}Var_t\left(m_{t+2}\right) + Cov(m_{t+1}, m_{t+2})$$

$$\begin{aligned} P_t^2 \quad = \quad & E_t\left(m_{t+1}\right) + E_t(\pi_{t+1}) + E_t\left(m_{t+2}\right) + E_t(\pi_{t+2}) + \\ & + \frac{1}{2}Var_t\left(m_{t+1}\right) + \frac{1}{2}Var_t(\pi_{t+1}) + \frac{1}{2}Var_t\left(m_{t+2}\right) + \frac{1}{2}Var_t(\pi_{t+2}) + \\ & + Cov_t(m_{t+1}, m_{t+2}) + Cov_t(\pi_{t+1}, \pi_{t+2}) + \\ & + Cov_t(m_{t+1}, \pi_{t+1}) + Cov_t(m_{t+1}, \pi_{t+2}) + \\ & + Cov_t(m_{t+2}, \pi_{t+1}) + Cov_t(m_{t+2}, \pi_{t+2}) \end{aligned}$$

The 2 period forward rates are given by:

$$f_t = p_t^1 - p_t^2$$

$$f_t^{\$} = -E_t\left(m_{t+2}\right) - \frac{1}{2}Var_t\left(m_{t+2}\right) - Cov(m_{t+1}, m_{t+2})$$

$$\begin{aligned} f_t \quad = \quad & -E_t\left(m_{t+2}\right) - E_t(\pi_{t+2}) - \frac{1}{2}Var_t\left(m_{t+2}\right) - \frac{1}{2}Var_t(\pi_{t+2}) - \\ & - Cov(m_{t+1}, m_{t+2}) - Cov(\pi_{t+1}, \pi_{t+2}) - \\ & - Cov_t(m_{t+1}, \pi_{t+2}) - Cov_t(m_{t+2}, \pi_{t+1}) - Cov_t(m_{t+2}, \pi_{t+2}) \end{aligned}$$

Current Account as a Dynamic Portfolio Choice Problem*

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May 20th, 2007

Abstract

The current account problem can be understood as the choice of where to allocate national savings: at home or abroad. Even being essentially a question of portfolio allocation, standard macroeconomic models did not incorporate this nature of the problem clearly until very recently. Commonly used models have static-like solutions with constant portfolios over time. In this paper, first we argue that the existing tests are powerless to reject the constant-share null under weak conditions. Once the test is modified, we find clear evidence against constant portfolio current account models. A thorough analysis of the data reveals that portfolio rebalance is indeed important. For this reason, we develop a current account model in which the representative agent's portfolio choice problem is a dynamic one. Time-varying investment opportunities is the main mechanism behind portfolio rebalances in our model. Thus, we are able to generate rebalancing in portfolios that in turn affects the current account. We estimate/solve this model using a long time series data from different assets in the US and Japan. We use the optimal fraction of the portfolio in foreign assets to explain the US-Japan bilateral current account. Results indicate that variations in investment opportunities can explain at least 54% of its movements, a performance superior to previous models.

JEL Classification: F11, G11, F37, F32

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1. Introduction

On the economic literature, traditional models do not provide empirical predictions for the current account that are matched by the data. For instance, what is the expected current account response to temporary income shocks? Until very recently, the answer was based on the neoclassical intertemporal approach to the current account.[1] In the simplest version of the model for a small open economy, the marginal unit of saving would be invested abroad. This in turn would lead to an improvement in the current account. Consumption smoothing and diminishing returns in the domestic economy are the mechanisms behind the model. In other words, the model suggests that the current account should respond one to one to changes in savings. However, the empirical evidence does not support this traditional rule in a variant of what is known as the Feldstein-Horioka Puzzle.

Kraay and Ventura (2000) modified this intertemporal approach by introducing investment risk as an important mechanism in determining how this marginal unit of saving should be allocated. Due to investment risk, this extra unit of saving should not necessarily be invested in a single asset. If the importance of diminishing returns in the domestic economy is relatively low in comparison to the importance of investment risk, part of this extra unit of savings should be invested at home. Then, the current account is fundamentally a portfolio choice problem. This was clearly a step forward relative to the old inter-temporal approach that did not take investment risk into account. However, they use Merton's (1971) model: the representative agent has log-utility, and returns and volatilities are constant over time. Under this formulation, agents behave as a mean-variance optimizers à la Markowitz-Tobin. In other words, the solution to the asset allocation problem is the same as if it were a static problem. Thus, the optimal portfolio can be found by looking at a country's portfolio today.[2] In order to maintain the optimal portfolio unchanged, the marginal unit of saving should be invested proportionally to investments in domestic and foreign assets. Kraay and Ventura (2000) propose, thus, this new rule: the marginal unit of saving resultant of a temporary income shock should be invested as the average unit. Interestingly, transitory income shocks lead to symmetrically opposite effects on the current account in creditor and debtor countries. A positive income shock leads to current account deficits in debtor countries and surpluses in creditor countries. The response of investment to a positive shock on a debtor country should be higher than a one to one increase in the savings. In other words, we should observe a leverage to invest at home inducing a current account deficit.

The claim is of strong empirical support for this "new rule". However, positive results are restricted to the between regressions, interpreted as long-run results. We argue that these long-run results could be generated by a large class of models. More precisely, we show that the between regression results are almost tautological under the weak conditions[3]. Therefore, these regressions do not shed light on the mechanisms of the model. This model can explain on average only 17% of current account movements, another piece of evidence against a constant portfolio rule.

The analysis should be made with time series regressions. However, the time series analysis does not support constant portfolio rules. The statistical properties of

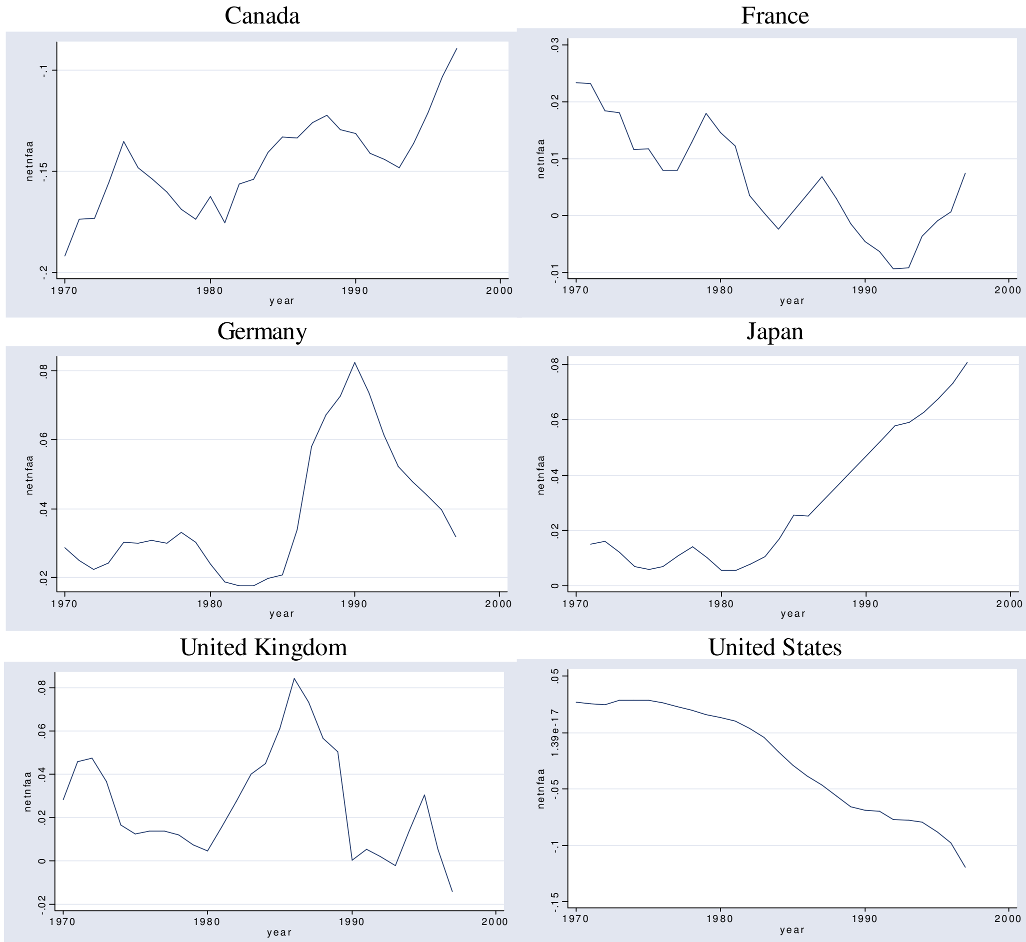
[1] See Obstfeld and Rogoff (1996).

[2] Specifically, one should look at the fraction of net foreign assets/total assets, as it will become clear later on.

[3] The conditions is low covariance between the ratio of net foreign assets to total assets and the savings rate

the countries' portfolio composition suggest rebalance away from constant composition. Figure 1 shows country's portfolio shares to foreign assets. As it can be seen, they are not constant over time. The persistence of changes in portfolio composition within countries indicates that portfolio rebalancing away from constant-shares is relevant in the short and long run.

Figure 1



There are many possible reasons for country portfolios not to be constant. We focus on one particular reason which we believe to be the most important among them: changes in investment opportunities.[4] We propose here an asset allocation model for the current account. Portfolio shares are time varying due to time varying investment opportunities. We assume agents with Epstein-Zin utility function and a VAR dynamic for asset returns and other state variables to obtain time-varying investment opportunities. Using Campbell, Chan and Viceira´s (2003) numerical solution method, we apply it to US and Japanese assets. The results are based solely on the time series of asset returns. We want to understand how well this model can explain current account dynamics.

The empirical tests of our model do provide strong evidence of this relevant mechanism behind current account dynamics. We show that, for the US and Japan,

[4] Other reasons could be time-varying preferences (risk aversion) and financial constraints.

variations in expected returns did change agents' optimal portfolio composition in a direction consistent with the actual bilateral current account movements. Variations in investment opportunities could be responsible for at least 54% of the pattern observed in the US bilateral current account with Japan. Therefore, we can conclude that time-varying investment opportunities are an important mechanism behind country portfolios and current account changes.

In sum, this paper first discusses theoretical specifications of the portfolio view of the current account and their empirical support. We argue that portfolio rebalance away from constant shares is indeed important. We propose a new current account model with dynamic portfolio allocation. The empirical evidence presented suggests that, at least for the bilateral current account of the countries analyzed (US and Japan), this model seems to provide superior empirical results.

Although we did not aim to draw direct policy implications, our results can also contribute to the current debate on global imbalances. Many competing explanations for global imbalances, specially the rising US current account deficit, were developed. One of them relies on the argument of better investment opportunities in the US compared to other G7 countries as Cooper (2004), Clarida (2005), and Backus and Lambert (2005). Our model can be understood as a formalization of this hypothesis. As shown by Lane and Milesi-Ferretti (2005), Japan is by far the largest net creditor in international investment positions and the US, the largest net debtor in international investment positions. The results indicate that a significant part (but not the total) of the US imbalances can be explained by better investment opportunities at home.

The rest of the paper is structured as follows. Section 2 derives theoretical portfolio approaches to the current account and discusses their different specifications. Section 3 discusses the constant portfolio model results and argues that portfolio rebalance away from constant shares is important. Section 4 presents our estimates of our model with time-varying portfolio shares for US and Japan. In this section, we also show that the results do explain significantly the dynamics of their bilateral current account. Section 5 concludes.

2. Theoretical Framework: a Portfolio Approach

The current account balance can be analyzed by two alternative focuses: on trade patterns (exports and imports) or on savings, investments, and international financial flows. Through an accounting identity, the current account balance in any given year must equal the excess domestic savings over domestic investment: $NX = CA = S - I$. This is the favorite theoretical framework for economists to analyze current account dynamics. Indeed, as highlighted in Bernanke (2005), it is the framework used nowadays by economists to find an explanation for the recent patterns in the US current account deficits:

"... specific trade-related factors cannot explain the magnitude of the U.S. current account imbalance and its recent sharp rise. Rather, the U.S. trade balance is the tail of the dog; for most part it has been passively determined by foreign and domestic incomes, asset prices, interest rate, and exchange rates, which are themselves in turn the product of more fundamental driving forces. Instead, an alternative perspective on current account appears likely to be more useful for explaining recent developments. This second perspective focuses on international financial flows and the basic fact that a country's saving and investment need not be equal each period." - Ben Bernanke (2005), page 2.

Under this second framework, changes in savings and investment decisions are directly reflected in the current account balance. There are basically two factors driving this relation: (i) income (wealth) shocks at home and abroad; and/or (ii) time-varying investment opportunities. The literature has traditionally focused on the former. Transitory income shocks induce current account responses due to a consumption smoothing force. The traditional intertemporal approach to the current account claims that this extra savings should be invested abroad, thereby improving the current account in a one to one basis. Kraay and Ventura (2000) stress that the resulting current account balances depend fundamentally on how the extra savings unit is invested. If part of this extra savings is invested at home, the current account response should be less than the in the former case. This suggests that the portfolio choice part of the problem cannot be ignored: how savings are invested is at least as important as the amount of savings generated. Under the existing theoretical formulations to understand the current account dynamics, the optimal portfolio is always constant over time. However, the empirical evidence shown in the next section does not support this claim. Thus, in the present section, we present a model in which the optimal portfolio might not be constant over time. The main mechanism behind time-varying portfolio shares in our model is time-varying investment opportunities (even without the presence of income shocks). The literature had not yet theoretically formalized this argument estimating its effect. Later, in section 4, we show that, empirically, this theoretical approach do provide a good explanation for the current account pattern observed between the US and Japan. Let us turn now to the model.

2.1. The General Model

The model is set in discrete time.[5] Consider a country populated by identical and infinitely lived individuals whose preferences can be represented by the time-separable utility function over their consumption stream.[6] The home country problem is the following:

$$\max_{C_t,\alpha_1,\alpha_2,\ldots,\alpha_1^*,\alpha_2^*,\ldots} U_t = u(C_t, E_t(C_{t+1}, C_{t+2}, \ldots)$$
$$s.t.\ \mathrm{W}_{t+1} = (W_t - C_t)R_{p,t+1} \tag{1}$$

where $u' > 0$, $u'' < 0$, δ is the time discount factor, W_t is the total wealth at time t, and $R_{p,t+1}$ is the country's portfolio gross return from t to $t+1$. Thus, equation (1) is the intertemporal budget constraint faced by the home country. If we assume time separable utility functions, $u(C_t) + E_t\sum_{j=1}^{\infty}\delta^j u(C_{t+j})$, the first order conditions for this maximization problem yield the famous Euler equation for consumption:

[5] We are considering here a partial equilibrium analysis. We solve the micro problem of an investor facing exogenous asset returns, but we do not argue how these asset returns could be consistent with a general equilibrium model.
[6] The hypothesis of infinitely lived individuals leads to identical results to the case where individuals care about their offspring, who, in turn, care about theirs.

$$E_t\left[\frac{u'(C_{t+1})}{u'(C_t)}\delta R_{p,t+1}\right]=1 \qquad (2)$$

The investor's optimal consumption and savings decisions must satisfy equation (2). When investment opportunities are constant, i.e. $R_{p,t+1}$ is time-invariant, the optimal policies imply a constant consumption-wealth ratio[7], and therefore a constant investment/wealth ratio. The same solution would be obtained if the problem was a static one.

Now, let's turn to the investment side of the problem. Agents can invest their money at home or abroad. There are *n* securities available for investment at home and *n* securities available abroad, so that *2n* is the total number of available securities. As examples of possible securities that agents could invest we can mention: home physical capital, foreign physical capital, home stocks, foreign stocks, money markets, bonds, etc. All income available for consumers at time *t* is generated by the returns of their previous period portfolio and by any sales they make (short sales of any asset are allowed). Obtaining a loan is equivalent to short the money market instrument. One can interpret a country's GDP as the total return of the investment in home assets.[8]

Define $\alpha_{i,t}$ as the proportion of home country's wealth invested in home asset *i* from *t* to *t+1* and $\alpha^*_{i,t}$ as the proportion of home country's wealth invested in foreign asset *i* from *t* to *t+1*. The home country's real portfolio return $R_{p,t+1}$ is then given by:

$$R_{p,t+1}=\sum_{i=1}^{n}\alpha_{\mathrm{i,t}}R_{i,t+1}+\sum_{i=1}^{n}\alpha^*_{i,t}R^*_{i,t+1} \qquad (3)$$

At each point in time, agents decide where and how much of their wealth will be allocated in each security. Thus, we can aggregate these investments to discover how much is allocated at home and abroad. The total wealth invested at home at time *t* is $\alpha_t W_t=\sum_{i=1}^{n}\alpha_{i,t}W_t$ and the total wealth invested abroad is $\alpha^*_t W_t=\sum_{i=1}^{n}\alpha^*_{i,t}W_t$. To simplify the analysis, also assume that foreigners cannot invest in domestic assets[9]. Under this framework, we can then use an accounting identity to define the current account balance CA_{t+1} as the variation in the wealth allocated abroad form *t* to *t+1*:

$$\begin{aligned}
&CA_{t+1}=\alpha^*_{t+1}W_{t+1}-\alpha^*_t W_t\\
&or, equivalently,\\
&CA_{t+1}=\alpha^*_t\Delta W_{t+1}+\Delta\alpha^*_{t+1}W_{t+1} \qquad (4)
\end{aligned}$$

Therefore, current account imbalances can be caused by income (wealth) shocks and by shocks to portfolio shares. These shares respond to time-varying investment

[7] If returns are IID, there are no changes over time in investment opportunities that might induce changes in consumption relative to wealth.
[8] Note that we do not allow for labor income.
[9] That is, we are presenting a partial equilibrium problem. Right now we are extending the model to a general equilibrium one.

opportunities, parameter uncertainty, time-varying risk aversion, among other reasons. The different models presented below for the current account will basically depend on different functional forms for the utility function *u(.)* and on different dynamics for the asset returns $R_{p,t+1}$.

2.2. Constant Portfolio Models

We refer to the traditional intertemporal approach to the current account and to the Kraay and Ventura (2000) models as constant portfolio models. The main feature of both models is the constant investment opportunities. Kraay and Ventura (2000) relies on Merton (1971)'s model where agents have logarithmic utility function ($u(C_t) = \log(C_t)$) and assets have a (constant) log-normal distribution. The log-utility assumption implies that the consumption will always be a constant fraction of wealth. Log-utility implies that the intertemporal elasticity of substitution is equal to one and therefore, substitution and income effects exactly cancel out.[10] This is the feature that allows for a closed form solution to the optimization problem above.

$$C_t^{Optimal} = bW_t$$

Where b is a constant.

In our discrete time version, the hypothesis that log of asset returns follow a Brownian motion can be represented by:

$$\log R_t = r_t \overset{i.i.d.}{\sim} N(\mu, \Sigma) \qquad \forall t$$

In this specification, the solution to the first order condition of how much to invest in each asset is given by:[11]

$$\underset{(2n,1)}{\vec{\alpha}} = \Sigma^{-1}(\mu - r_f i + \sigma^2 / 2) \tag{5}$$

where σ^2 is the vector containing the diagonal elements of Σ and i is a $2n$-dimension vector of ones.

This result is equivalent to the Markowitz mean-variance model. Given agents preferences and asset's risk return characteristics, there is an optimal portfolio rule that, in turn, depends on assets risk and return. The distribution of asset returns is constant over time. Thus, there is no need to change portfolio shares: $\alpha_{t+1}^* = \alpha_t^* = \alpha^*, \forall t$. Hence, at any given instant, agents already hold their optimal portfolios, $\alpha^* = NFA/W$, where NFA is net foreign assets and W is total assets (wealth).

This model has clear cut predictions to the current account. In order to maintain its (optimal) portfolio unchanged, the marginal unit of savings in any given year should be invested as the average savings is invested. Average unit of saving is characterized by the yesterday's portfolio and the optimal fraction to be allocated abroad

[10] A log-utility function is a power utility (also known as CRRA) with risk aversion coefficient equals to one: $u(C_t) = \dfrac{C_t^{1-\gamma}}{1-\gamma}$. Under this formulation, the intertemporal elasticity of substitution is the reciprocal of the risk aversion coefficient. This is the main difficulty of this specification since there is no particular reason for these two coefficients to be related in any way.

[11] The proof is in many textbooks as in Campbell et al (2000).

corresponds to the fraction of country's net foreign assets over total assets (wealth). This implies that the current account response in net debtor and net creditor countries are the opposite of each other. Positive income shocks lead to current account deficits in debtor countries and current account surpluses in creditor countries. The intuition behind the net debtor country response is the following: in response to a positive shock, the extra unit of savings is invested at home and, in order to maintain the proportional leverage, the country should take even more debt. Hence, this rule implies that the Feldstein-Horioka finding is simply a flow version of the home bias in country portfolios[12].

Using our current notation, this rule is represented by: $CA_t = \alpha^* \Delta W_t = \alpha^* S_t = (NFA/A) S_t$. This equation is used by Kraay and Ventura (2000) to test their "new rule".

This was a clear improvement over traditional intertemporal models, which did not even take into account investment risk. In this framework, the prediction of the traditional intertemporal current account model would be observed if diminishing returns in the domestic economy was sufficiently important. In other words, if the marginal unit of savings invested at home caused returns on domestic assets to diminish abruptly. In this scenario, marginal savings should all be invested abroad and, therefore, the current account would improve one for one with a transitory income shock.[13] In our notation above, this means that $\alpha^* = 1$. Hence, according to this standard model: $CA_t = \Delta W_t = S_t$. This is the core implication of this model, traditionally used to test the intertemporal approach for the current account. It also leads us to a variant of what is known as the Feldstein-Horioka Puzzle. We will explore this empirical test in the next section.

Note that both approaches produce optimal portfolio constants over time. In the "traditional rule" all of the extra savings are invested abroad and, thereby, the current account always improves one to one with income shocks. In Kraay and Ventura's "new rule", the investment response, and therefore the current account dynamics, depends on the current portfolio.

2.3. A Dynamic Portfolio Model

To generate a dynamic portfolio choice, we choose to relax the hypothesis that investment opportunities are constant over time. We allow them to follow a first-order VAR. If markets were efficient, this specification would necessarily be reduced to a random walk, as we would not find any significance on any of the state variables in this VAR. However, the financial economics and behavioral finance literatures have been pointing out that financial returns are predictable to some extent.[14] Variables like past returns, earnings-to-price ratio, and the term spread have been documented to predict returns on many asset classes. Later, we will use these state variables on the estimation of our VAR. For now, let this be represented by a general vector z_{t+1} containing the asset returns and the state variables (s_{t+1}) follow a VAR dynamic:

[12] If we have home equity bias, according to this model, a significant fraction of the country's savings will be invested at home to maintain this current (optimal) portfolio. Therefore, we should observe a low correlation between savings and the current account or, equivalently, a high correlation between savings and investment, what characterizes the Feldstein-Horioka Puzzle.

[13] Kraay and Ventura (2000) give a mathematical proof for this statement.

[14] See De Bondt and Thaler (1985), Jagadeesh and Titman (1993), Fama and French (1992), Shleifer (2000), among many others.

$$z_{t+1} = \Phi_0 + \Phi_1 z_t + v_{t+1} \tag{6}$$

Where

$$v_{t+1} \overset{i.i.d.}{\sim} N(0, \Sigma_v)$$

$$z_{t+1} = \begin{bmatrix} r_{1,t+1} \\ x_{t+1} \\ s_{t+1} \end{bmatrix}, x_{t+1} = \begin{bmatrix} r_{2,t+1} - r_{1,t+1} \\ r_{3,t+1} - r_{1,t+1} \\ \dots \end{bmatrix}$$

Even with time-varying investment opportunities, if investors had a log-utility function, the savings decision would be reduced to a static one[15]. To be as general as possible, we allow agents to have Epstein-Zin (1989, 1991) recursive preferences. This functional form for the utility function has the advantage of separating the elasticity of intertemporal substitution from the relative risk aversion parameter. As already argued, there is no particular reason why these two parameters should be related to each other. This utility function nests as a special case the power utility specification, in which one is the reciprocal of the other, and the log-utility specification, in which both are equal to one. Epstein-Zin utility function is recursively defined by:

$$U(C_t, E_t(U_{t+1})) = \left[(1-\delta) C_t^{\frac{1-\gamma}{\theta}} + \delta \left(E_t \left(U_{t+1}^{1-\gamma} \right) \right)^{\frac{1}{\theta}} \right]^{\frac{\theta}{1-\gamma}}$$

where

$$\theta \equiv \frac{1-\gamma}{1-\psi^{-1}}$$

Here, $\gamma > 0$ is the relative risk aversion coefficient and $\psi > 0$ is the elasticity of intertemporal substitution. The first order condition for the problem stated in (1) assuming these preferences was solved by Epstein-Zin (1989):

$$E_t \left[\left\{ \delta \left(\frac{C_{t+1}}{C_t} \right)^{-\frac{1}{\psi}} \right\}^{\theta} R_{p,t+1}^{-(1-\theta)} R_{i,t+1} \right] = 1 \qquad for\ \forall i \tag{7}$$

To make predictions for the current account, we need to solve the model and find the time series of α_t^* generated. There is no closed-form solution for this problem. Fortunately, Campbell, Chan, and Viceira (2003) have recently proposed an approximate solution method for this multivariate model of strategic asset allocation. They showed that this problem can be approximately reduced to a system of linear quadratic equations for portfolio weights and consumption as affine functions of the state variables.

Let's start with equation (3), the portfolio return equation. It can be re-written in the following way:

[15] See Campbell and Viceira (2002) for an extensive study on strategic asset allocation.

$$R_{p,t+1}=\sum_{i=2}^{n}\alpha_{i,t}(R_{i,t+1}-R_{1,t+1})+\sum_{i=1}^{n}\alpha^{*}_{i,t}(R^{*}_{i,t+1}-R_{1,t+1})+R_{1,t+1} \tag{8}$$

Where the first asset, whose real return is given by $R_{1,t+1}$, is a domestic short-term instrument used as a benchmark asset. Even though all other returns are measured relative to this benchmark asset, it is not assumed to be riskless.

The log return can then be approximated as:

$$r_{p,t+1}\approx r_{1,t+1}+\alpha_t' x_{t+1}+\frac{1}{2}\alpha_t'(\sigma_x^2-\Sigma_{xx}\,\alpha_t)+\alpha_t^{*}{}' x^{*}_{t+1}+\frac{1}{2}\alpha_t^{*}{}'(\sigma_x^{*2}-\Sigma^{*}_{xx}\,\alpha^{*}_t) \tag{9}$$

Where $\sigma_x^2\equiv diag(\Sigma_{xx})$, similarly defined for the foreign assets, is the variance of the excess returns. This approximation holds exactly in continuous time and it is highly accurate for short-time intervals.

The next equation is the budget constraint, equation (1). Log-linearizing it around the unconditional mean of the log-consumption-wealth ratio to obtain:

$$\Delta w_{t+1}\approx r_{p,t+1}+\left(1-\frac{1}{\rho}\right)(c_t-w_t)+k \tag{10}$$

where

$$\rho\equiv 1-\exp(E[c_t-w_t])$$

$$k\equiv\log(\rho)+(1-\rho)\log(1-\rho)/\rho$$

This form of the budget constraint holds exactly if the elasticity of intertemporal substitution $(\psi=1)$ is equal to 1, in which case $\rho=\delta$ and c_t-w_t is constant.

Finally, a second-order Taylor expansion to the Euler equation (7) around the conditional means of $\Delta c_{t+1}, r_{p,t+1}, r_{i,t+1}$ to obtain:

$$\theta\log\delta-\frac{\theta}{\psi}E_t\Delta c_{t+1}-(1-\theta)E_t r_{p,t+1}+E_t r_{i,t+1} \\ +\frac{1}{2}\mathrm{var}_t\left[-\frac{\theta}{\psi}\Delta c_{t+1}-(1-\theta)r_{p,t+1}+r_{i,t+1}\right]\approx 0 \tag{11}$$

This form of the Euler equation holds exactly if consumption and asset returns are jointly log-normally distributed, which is the case when $\psi=1$.

Thus, the model can be approximately solved based on these three equations. Campbell, Chan, and Viceira (2003) show that the solution is characterized by the following optimal portfolio weights and consumption rules:

$$\alpha_t = A_0 + A_1 z_t \tag{12}$$

Where

$$A_0 = \left(\frac{1}{\gamma}\right)\Sigma_{xx}^{-1}\left(H_x\Phi_0 + \frac{1}{2}\sigma_x^2 + (1-\gamma)\sigma_{1x}\right) + \left(1-\frac{1}{\gamma}\right)\Sigma_{xx}^{-1}\left(\frac{-\Lambda_0}{1-\psi}\right) \tag{13}$$

$$A_1 = \left(\frac{1}{\gamma}\right)\Sigma_{xx}^{-1} H_x\Phi_1 + \left(1-\frac{1}{\gamma}\right)\Sigma_{xx}^{-1}\left(\frac{-\Lambda_1}{1-\psi}\right) \tag{14}$$

Λ_0 and Λ_1 are constants.

$$c_t - w_t = -\rho\psi\log\delta - \rho\chi_{p,t} + \rho(1-\psi)E_t(r_{p,t+1}) + \rho k + \rho E_t(c_{t+1} - w_{t+1}) \tag{15}$$

Where

$\mathrm{E}_t(r_{p,t+1})$ and $\chi_{p,t}$ are quadratic functions of the VAR state variables.

This model has multiple state variables, and therefore it is not feasible to obtain an approximate closed-form solution. A numerical procedure will therefore be used to solve for the consumption and portfolio choice given ρ.

This is a general theoretical framework for the strategic asset allocation problem we are interested in. The model yields optimal portfolio rules that are linear in the vector of state variables, and therefore, a time-varying optimal allocation. In section 4, we will come back to the empirics of this version of the model. We will analyze to what extent this model can account for the current account dynamics. More specifically, we will apply it to investors allocating their portfolios among several different assets in the US and in Japan. It is then possible to construct a time series of α_t^* and compare it with the time series of the bilateral current account between these two countries.[16] But first, let's take a close look at the data and test if is it really necessary to have a dynamic portfolio model to explain current account.

3. Empirical Tests of the Current Account Theory

In this section, we use data for 13 industrial countries from 1970 to 1998 to perform empirical tests of the current account models.[17] We focus on industrial countries to avoid dealing with problems inherent to developing countries such as sudden stops. Furthermore, in the next section we will explore even further the bilateral current account of the two biggest economies during this period, the US and Japan.

We start by replicating traditional tests with our sample to clarify the discussion and show that our empirical criticisms are not explained by a different data set. First, we test constant portfolio models. The traditional intertemporal current account model prediction is not met by the data in a variant of the famous Feldstein-Horioka Puzzle. On the other hand, tests of the Kraay and Ventura (2000) model have impressive results on the cross-sectional dimension, but the time series results are not good. We discuss these results in some depth and conclude that one should not rely on the cross-

[16] In our approach, as in the other constant portfolio current account models presented before, home and foreign assets are treated as perfect substitutes. In contrast, Blanchard, Giavazzi, and Sa (2005) treat them as imperfect substitutes.

[17] The countries included in the analysis are: Australia, Austria, Canada, Finland, France, Germany, Italy, Netherlands, Japan, Spain, Sweden, US, and UK.

sectional results. It is demonstrated that between regression would yield the desired coefficient estimate under weak conditions. Not only the rejection of these constant portfolio models, but also the time series regressions and the autoregressive properties of country portfolios lead us to the conclusion that portfolio rebalance away from constant shares is important to explain current account dynamics.

Concerning data sources, the data for Net Foreign Assets is from Lane and Milesi-Ferretti (2004). They have constructed a dataset that considers capital gains/losses on country portfolios. The data for Total Wealth is from Kraay, Loayza, Serven and Ventura (2000). Since this variable is not available after 1998, our sample for the purposes of this section ends in 1998 as well. Current account and national savings data are from IMF's International Financial Statistics. The Solow residual and population growth data are the same as Kraay and Ventura (2000).

3.1. Testing Constant Portfolio Models

Tests for both the traditional intertemporal current account rule and for Kraay and Ventura (2000) new rule are based on a simple regression which tests one of the model's predictions. For the traditional rule, the empirical implication tested is that extra savings should all be invested abroad. Formally the test is:

$$CA_{i,t} = \alpha + \beta S_{i,t} + \varepsilon_{i,t,} \qquad (16)$$
$$\mathrm{H}_0 : \beta = 1$$

The subscript *(i,t)* denotes time *t* and country *i*. Note that since $CA_t = S_t - I_t$, testing equation (16) is the same as the famous Feldstein-Horioka regression $S_{i,t} = \delta + \gamma I_{i,t} + \upsilon_{i,t}$. In other words, testing β=1 is equivalent to testing γ=0. Similarly, Kraay and Ventura's new rule tests are done over its empirical prediction that marginal savings should be invested as the average savings:

$$CA_{i,t} = \alpha + \beta \frac{NFA_t}{W_t} S_{i,t} + \varepsilon_{i,t,} \qquad (17)$$
$$\mathrm{H}_0 : \beta = 1$$

Where NFA_t is the home country's net foreign assets, which in turn is equal to $\alpha_t^* W_t$.

Tables 1 and 2 summarize the results of these tests with our dataset. We performed three types of estimations: (i) pooled regressions; (ii) between regressions, cross-sectional regression on the time series average for each country; and (iii) within regressions, the usual fixed effects panel-data regression. For all of them, we do standard least squares regressions and instrumental variables regressions, in which ranks of the right hand side variables are used as instruments to correct for any measurement error. As it is usually done in this literature, we also included the Solow residual and population growth as controls in the current account regressions.[18]

[18] Kraay and Ventura (2000) also include these controls in their empirical tests.

The tests for the traditional rule are presented in Table 1. The coefficient β varies from 0.180 to 0.275. The null hypothesis that β is equal to 1 is always rejected. This is equivalent to the famous Feldstein-Horioka (1980) puzzle.[19]

Table 1
Tests of the Traditional Rule

	Pooled Regression		Between Regression		Within Regression	
	OLS	IV	OLS	IV	OLS	IV
Gross National Savings/GNP	0.264	0.265	0.134	0.124	0.421	0.405
	[9.651]***	[7.513]***	[0.883]	[0.779]	[6.477]***	[5.518]***
Solow Residual	0.149	0.150	-1.709	-1.764	0.191	0.193
	[1.439]	[1.439]	[1.052]	[1.069]	[2.750]***	[2.770]***
Population Growth	-1.355	-1.354	-2.822	-2.856	0.145	0.155
	[5.127]***	[5.149]***	[2.082]*	[2.088]**	[0.373]	[0.395]
Constant	-0.070	-0.071	-0.011	-0.008	-0.109	-0.090
	[5.753]***	[5.341]***	[0.252]	[0.183]	[5.743]***	[4.426]***
Country Dummies	No	No	No	No	Yes	Yes
Year Dummies	Yes	Yes	No	No	Yes	Yes
No. of Obs.	247	247	247	247	247	247
R-squared	0.38	0.38	0.495	0.495	0.7	0.7

Robust t statistics in brackets
* significant at 10%; ** significant at 5%; *** significant at 1%

Table 2 reports the test results of the new rule for the current account response to temporary income shocks. Pooled regressions provide some evidence that the current account moves more than one to one with changes in savings, contradicting the theory in test. This means that lump sum income shocks lead countries to change their portfolio composition: debtors increase their net debt and creditors invest even more abroad. However, as in Kraay and Ventura (2000), the within and between regressions seem to support the new rule, with a β not significantly different than 1. The between regressions reflect cross-sectional variation and thus, predict a stronger relationship than the within regressions.

[19] They found that the coefficient of the regression of Investments on Savings is very high, being really close to 1. The result that savings and investments are strongly correlated indicates that international financial markets are not perfectly integrated, as it is well documented in the International Finance literature.

Table 2 Tests of the New Rule						
	Pooled Regression		Between Regression		Within Regression	
	OLS	IV	OLS	IV	OLS	IV
Gross National Savings/GNP X (Foreign Assets/Total Assets)	1.341	1.328	1.369	1.328	0.988	1.001
	[12.756]***	[12.356]***	[5.672]***	[5.444]***	[3.923]***	[3.672]***
Solow Residual	0.199	0.198	-0.385	-0.449	0.243	0.243
	[2.349]**	[2.343]**	[0.506]	[0.587]	[3.186]***	[3.184]***
Population Growth	-0.126	-0.142	-0.539	-0.622	0.352	0.352
	[0.500]	[0.567]	[0.694]	[0.798]	[0.917]	[0.917]
Constant	-0.013	-0.013	0.005	0.006	-0.007	0.003
	[1.319]	[1.315]	[0.726]	[0.810]	[0.691]	[0.258]
Country Dummies	No	No	No	No	Yes	Yes
Year Dummies	Yes	Yes	No	No	Yes	Yes
No. of Obs.	247	247	247	247	247	247
R-squared	0.582	0.582	0.88	0.88	0.663	0.663
t-test of beta = 1	0.001	0.003	0.160	0.179	0.963	0.997

Robust t statistics in brackets
* significant at 10%; ** significant at 5%; *** significant at 1%

3.2. A Critical View of This Evidence

So far, what can we say about constant portfolio models? First, as in many other papers, the evidence points towards the failure of the traditional intertemporal current account rule. On the other hand, the results seem supportive of the new rule of a constant portfolio model in the cross-sectional dimension. The between regressions results can be interpreted as capturing long-run dynamics accurately. It uses the information on the time-series average (over 20 years or so) of country variables. Nevertheless, the relevant coefficient, β, is close to the expected[20] one, surprisingly so given the small sample considered. In this section, we argue that the between regression is almost tautologically bound to give the above results under very weak conditions. In other words, it can be explained by a wide class of models. Kraay and Ventura's (2000) model is one possible approach out of many.

Let us consider a model with a steady state optimal portfolio allocation. To simplify the exposition, suppose that there are only two kinds of assets: home domestic capital k and foreign capital k^*. Let us assume that foreigners cannot hold domestic capital. Total wealth is then given by $W_t = k_t + k_t^*$. Also, assume that output and capital grow at a constant rate g. Therefore:

$$\begin{aligned} S_t &= \Delta W_t \\ &= \Delta k_t^* + \Delta k_t \\ &= g(k_t^* + k_t) \end{aligned}$$

Moreover, on the steady state, investment must satisfy:

$$I_t = \Delta k_t = g k_t$$

[20] In the next section, we will provide some evidence that the time series results do not seem to support this new rule.

Hence, we can use the accounting identity again to obtain an equilibrium condition for the current account:

$$
\begin{aligned}
CA_t &= S_t - I_t = gk_t^* \\
&= \frac{k_t^*}{k_t^* + k_t} g(k_t^* + k_t) \\
&= \frac{NFA_t}{W_t} S_t
\end{aligned}
\tag{18}
$$

This argument is adapted from van Wincoop (2002). He argues that Kraay and Ventura's main contribution is to propose that NFA_t/W_t as a constant. We go one step further. Without imposing a constant NFA_t/W_t, it is still possible to obtain the between regression results assuming very little else.

Equation (18) implies the following:

$$1/T \sum_t CA_t = 1/T \sum_t (NFA_t/W_t) S_t \tag{19}$$

Let this be denoted:

$$\overline{CA} = \beta \overline{S_t (NFA_t / W_t)} \tag{20}$$

$$\textit{Where } \beta = 1$$

We also know that the between regression is given by the following:

$$\overline{CA_i} = \alpha + \beta \overline{\mathrm{S}}_{\mathrm{i}} \left(\overline{NFA_i/W_i} \right) + \varepsilon_{\mathrm{i}} \tag{21}$$

Under what conditions does equation (20) imply that β=1 in equation (21)? Equation (21) will be true if, for each country, the covariance between the net foreign assets to total assets ratio and the savings rate is low. The reason is as follows:

$$
\begin{aligned}
\overline{S \frac{NFA}{W}} &= \overline{\frac{NFA}{W}} \overline{S} + \mathrm{cov}(\frac{NFA}{W}, S) \\
&= \overline{\frac{NFA}{W}} \overline{S} + \rho \sigma_{NFA/W} \sigma_S
\end{aligned}
\tag{22}
$$

Therefore, equation (20) and (21) would be equivalent (so that β=1 will follow automatically) if the last term in equation (22) is negligible. This last term is small if the standard deviations of the net foreign assets to total assets ratio and the savings rate are low in absolute values and/or if the correlation coefficient ρ is low in absolute value. In fact, these conditions are satisfied for most countries. And indeed, savings rate are very low volatility are very low. Table 3 shows the relationship between the components of equation (22) for each country. For only 2 countries in our sample, the mean of the product and the product of the means differ by more than 10% (these are the US and Italy). The empirical evidence shows that on average the condition for the between regression is satisfied.

Thus, as long as a theoretical model has a steady state optimal portfolio allocation, and as long as it accepts that the standard deviations of x_t and S_t over time are small, we will have β=1. Therefore, the null hypothesis should not be rejected for the between regression. If the actual world portfolio allocation in fact has a steady state and the standard deviations are small (which seems to be the case), the null of β=1 will in fact not be rejected by the data. Even further, if the correlation ρ is positive, then the estimated β will exceed 1. This means that a very large class of theoretical models would predict the same result as the new rule for the between regression. Therefore, we need other tests on the mechanisms behind the current account dynamics in order to argue that a constant portfolio model is an accurate one[21].

Table 3
Correlations

Country	(xS)bar	xbar*Sbar	% deviation
Australia	-0.023	-0.024	3.443
Austria	-0.010	-0.010	1.832
Canada	-0.020	-0.020	0.147
Finland	-0.022	-0.023	2.915
France	-0.001	-0.001	1.082
Germany	0.007	0.007	2.898
Italy	0.000	0.000	167.935
Japan	0.011	0.011	1.198
Netherlands	0.019	0.019	0.432
Spain	-0.011	-0.011	0.318
Sweden	-0.009	-0.010	3.769
UK	0.004	0.004	1.190
US	-0.002	-0.003	21.961

Indeed, when we break the cross-sectional evidence in a year-by-year basis, so that our criticism is no longer value because we are not using between regression, we cannot accept that β=1 as reported in Table 4.

Table 4
Cross-Sectional Estimates of Constant Portfolio Models

	(S/Y)*(NFA/W)		**t-test** (p-value)		
Year	**Coef.**	**t-Stats**	**beta = 1**	**No. Obs.**	**R-Squared**
1975	1.811	[6.301]***	0.000	9	0.834
1976	1.269	[5.219]***	0.001	9	0.701
1977	1.000	[1.801]	0.115	9	0.232
1978	1.018	[1.335]	0.224	9	0.216
1979	0.848	[2.122]*	0.067	10	0.425
1980	0.956	[2.547]**	0.034	10	0.395
1981	1.826	[3.063]**	0.016	10	0.557
1982	1.363	[2.133]*	0.059	12	0.310
1983	1.593	[5.128]***	0.000	12	0.701
1984	1.540	[2.999]**	0.013	12	0.473
1985	1.967	[4.575]***	0.001	12	0.600
1986	1.913	[3.369]***	0.007	12	0.514
1987	1.450	[3.721]***	0.004	12	0.567
1988	1.292	[3.155]**	0.010	12	0.571
1989	1.716	[4.162]***	0.002	13	0.693
1990	1.605	[11.416]***	0.000	12	0.780
1991	1.586	[7.984]***	0.000	12	0.787
1992	1.680	[8.309]***	0.000	12	0.811
1993	1.518	[6.072]***	0.000	12	0.773
1994	1.235	[3.243]***	0.009	12	0.615
1995	0.815	[1.495]	0.166	12	0.289
Average	1.429				
Std. Dev.	0.350				

3.3. Testing Portfolio Rebalance

[21] It must be noted that the result of beta equals to one holds in the Within (Panel Data) estimation. However, since it is imposed that the betas must be the same in this specification, the coefficients would be affected anyways by the "curse" of low covariance of S and NFA/A. So, we believe that our criticism in the result remains.

At this point, it is clear that we cannot rely only on the between regressions since, as already argued, positive results for constant portfolio would follow under weak conditions. Thus, the natural way to start is to explore the time series evidence. Relaxing the restriction that β is constant across countries gives us more detailed information about the relationship between current accounts and portfolio choice. Table 5 reports the results. The time series evidence does not point toward a constant β across countries, nor towards a β = 1.[22] Moreover, there is wide dispersion away from β = 1, implying that the time series variation is not supportive of a constant portfolio rule for current account.

Table 5
Time Series Estimates

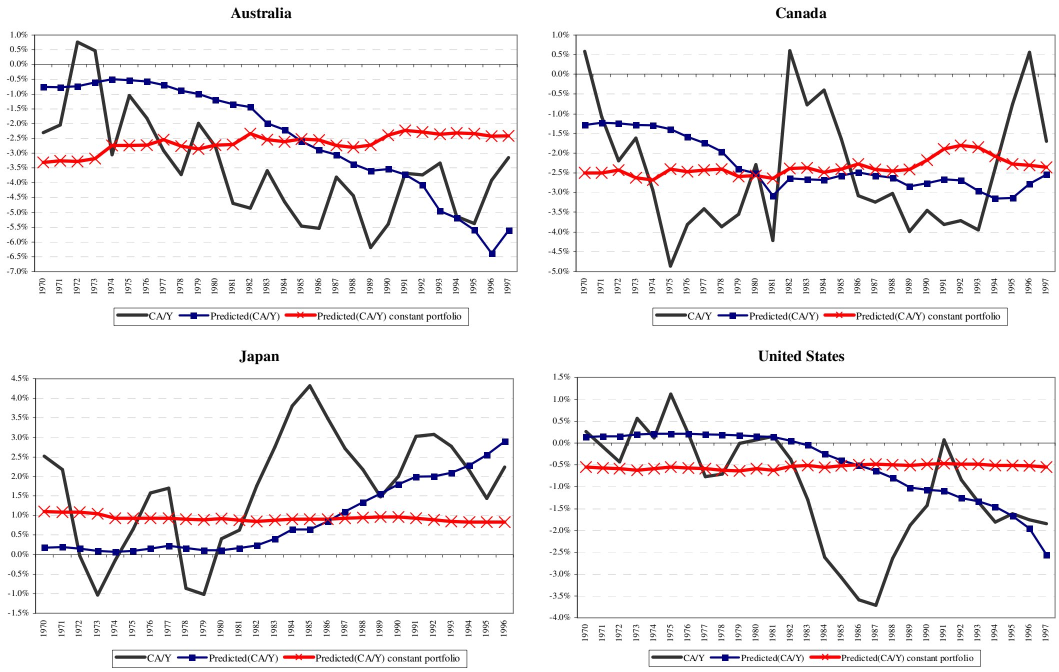
Country	(S/Y)*(NFA/W) Coef.	t-Stats	t-test (p-value) beta = 1	No. Obs.	R-Squared
Australia	1.054	[2.506]**	0.020	23	0.181
Austria	1.283	[1.171]	0.255	23	0.066
Canada	0.911	[1.092]	0.287	23	0.041
Finland	0.045	[0.031]	0.975	21	0.001
France	1.300	[0.304]	0.773	7	.
Germany	3.510	[6.681]***	0.000	15	0.551
Italy	-0.080	[0.064]	0.949	23	.
Japan	1.369	[3.455]***	0.004	17	0.192
Netherlands	0.698	[0.758]	0.463	14	0.083
Spain	3.562	[2.698]**	0.014	21	0.287
Sweden	-1.206	[0.666]	0.518	14	0.098
UK	-0.201	[0.187]	0.853	23	0.008
US	1.631	[4.501]***	0.000	23	0.364
Average	1.067				
Std. Dev.	1.354				

Robust t statistics in brackets
* significant at 10%; ** significant at 5%; *** significant at 1%

Figure 2 plots the time series of the observed Current Account/GDP (in solid black line) and the predicted ones by two different specifications: $S_t(NFA_t/W_t)$ (blue with squares), representing the "constant portfolio" rule (but since NFA/W is varying over time, it is not a truly constant portfolio model prediction); and the $S_t(NFA/W)$ (red with cross line), where we force the portfolio to be constant, being thus, a true constant portfolio rule. Since NFA_t/W_t varies over time, one can argue that this is not the "true prediction" of a constant portfolio theory. For this reason, we added the second one. The fitting of the figures are not impressive. The percentage of current account movements explained by this constant portfolio rule is very small[23]. Then, it is possible to state that, for an average country, a constant portfolio rule can explain at most 17% of the movements in its current account. We will explore the cases of the US and Japan in section 4, so it is worth noting that their R-squared are respectively 0.36 and 0.19.

[22] This is a restriction imposed by the within regression.
[23] We can interpret the R-squared of these regressions as the % of the current account movements in each country explained by the time series of NFA_t/W_t .

Figure 2



Not only the percentage of the current account movements explained by a constant portfolio rule is small based on the time series evidence, but also, in practice, the portfolio composition do seem to vary. The predicted CA_t in the graphs above is varying not because S_t is varying, but because NFA_t / W_t is varying. Based on the analysis of the time-series properties of NFA/W we can argue that, indeed, portfolio composition do vary over time. Table 6 reports the AR(1) coefficients and the implied half-lives of portfolio composition shocks: 86% of the 21 OECD countries analyzed displayed an AR(1) coefficient higher than 0.85, which implies a half-life longer than 4.3 years for these countries.[24] This is a highly persistent figure to be consistent with constant portfolios.

[24] Half-life is the amount of time needed for 50% of the effect of a given shock to dissipate.

Table 6
Half-Life of Portfolio Composition (NFA/W)

Country	AR(1)	Half-Life	No. Obs.
Australia	1.029	∞	27
Austria	0.908	7.166	27
Belgium	0.735	2.248	17
Canada	0.928	9.276	27
Denmark	0.853	4.347	18
Finland	0.960	17.155	27
France	0.860	4.582	27
Germany	1.178	∞	27
Greece	0.965	19.512	18
Italy	0.914	7.708	19
Japan	1.100	∞	14
Netherlands	1.105	∞	22
New Zealand	0.907	7.101	19
Norway	1.080	∞	24
Portugal	0.774	2.708	19
Spain	0.951	13.682	27
Sweden	0.898	6.416	18
Switzerland	0.892	6.089	13
Turkey	0.796	3.035	19
UK	0.934	10.104	27
US	1.090	∞	19
Average		8.075	

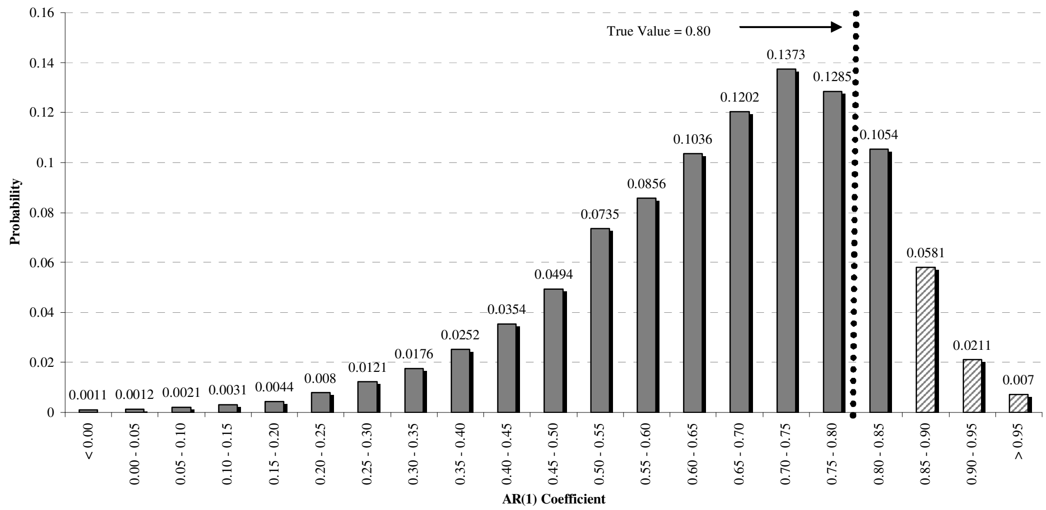
Surprising though were the results of an AR(1) coefficient greater than 1. We could interpret this evidence as a continuous shift of portfolios in the last 20 years. The US is a typical example of this case. We are also aware of the small sample size (around 27 observations for each country) and the possibility of a small sample bias. However, this would be a downward bias. Thus, incorporating this possibility into the previous analysis would yield parameters even more favorable to our hypothesis. In order to control for the downward small sample bias, we run 10,000 simulations of 27 observations each with a true AR(1) parameter value of 0.8 (half-life of 3.1 years). This choice for the true AR(1) parameter is arbitrary, but as argued below, different parameter choices would not weaken the argument made here.

The simulation results show that at a 10% significance level, the critical value for rejection that beta equals to 0.8 is 0.85. The histogram of the distribution of the AR(1) coefficients can be seen in Figure 2. This implies that in our sample, at 10% significance level, 85% of the countries have a half-live higher than 3.1 years (beta = 0.8). On a more conservative case, choosing a true beta of 0.87, we find from our simulations that at a 5% significance level, 52% of the countries in our sample have a half-life higher than 5 years. Ventura (2003) recons the time series failure of the model and tries to reconcile the short-run failure of the model by introducing adjustment costs. However, based on the very long half-lives of shocks, we can argue that in the short run, the adjustment cost story alone does not seem to explain the observed rebalances. If one accepts that 5 years to readjust 50% of a deviation from optimal portfolio is quite a long time bur reasonable, then the adjustment cost story could explain at most 50% of the cases in OECD countries.

Figure 3

Distribuition of AR(1) Coefficient on Samall Samples

10,000 simulations of 27 observations each



Let us now summarize the main points of this section. First, time series regressions do not seem to support a constant portfolio rule since most of the β coefficients are statistically different from one. Moreover, the percentage of current account movements explained by constant portfolio rules is on average low (at most 17%). It should also be noted that this is the case under the best possible scenario, since we let NFA_t / W_t vary over time, contradicting the constant portfolio theory. By investigating the statistical properties of the country portfolio compositions, there is evidence of large persistence of shocks in most of the OECD countries analyzed. To conclude, the empirical evidence suggests that there is persistence in portfolio rebalance away from constant shares on country portfolios. We now turn to one of the possible explanations for the portfolio rebalance: time varying investment opportunities.

4. A Dynamic Current Account Model with Time-Varying Investment Opportunities: an application to USA-Japan

As already argued, constant portfolio models do not seem to provide a proper explanation for the time series of current account dynamics. The empirical evidence suggests that country portfolios do vary over time, contradicting the most important hypothesis of constant portfolio models. In this section, we explore how well can time varying investment opportunities account for the observed movements of the US-Japan bilateral current account.

The current account balance depends on the portfolio choice made over the amount of national savings. Time-varying investment opportunities could, in principle, induce changes in the optimal portfolio which, in turn, would lead to capital flows reflected in the current account. We focus on this mechanism to explain current account movements. It is worth noting that this is not the only possible explanation. There are other possible reasons such as time varying preferences, parameter uncertainty, or financial constraints (e.g. if default is possible). However, we do not attempt to cover these or any other possibilities in this paper.[25]

Our approach consists in estimating and solving the model described in Section 2.3, a general theoretical framework for strategic asset allocation. This specification of the model has two important and interesting features: (i) it allows asset returns to follow a first-order VAR with state variables known in the empirical finance literature to predict asset returns; and (ii) the representative agent have an Epstein-Zin utility function, which allows us to model savings decisions and does not impose any link between the parameters for risk aversion and elasticity of intertemporal of substitution.

We implement this exercise for several assets from Japan and the US and then analyze the time series of the model's optimal proportion of US wealth that should be allocated to Japanese assets. We also consider the model's optimal proportion of Japan's wealth that should be allocated to US assets. Our results indicate that investment opportunities can explain more than 54% of the actually observed movements in the bilateral current account. This is an interesting result, especially if one notice that in the estimation of the model we do not use any information concerning the observed current account.

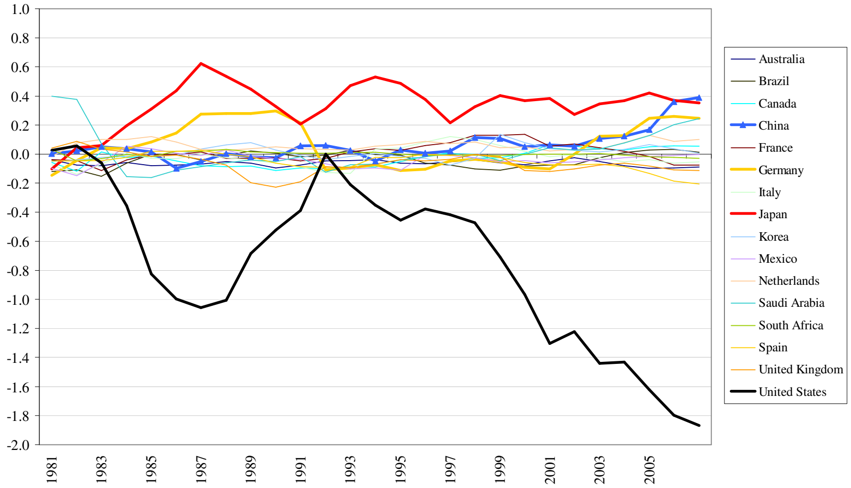
There are many reasons for choosing the US and Japan as the countries in our exercise. An important reason for choosing them is data availability. In order to estimate the model, we need a long time series of the asset returns. These data were available in a quarterly frequency for the US and Japan since 1960. The same is true for the US-Japan bilateral current account data. In our exercise, we use data on stock market returns, interest rates, and private firm profits (return on equity) for both the US and Japan.

But the main reason for choosing the USA and Japan is their economic relevance for international capital flows from 1960 to 2005 making them vary representative for analyzing the so-called “global imbalances”[26]. Figure 4 plots the data. The current account deficit in the US has been soaring and reached 7% of its GDP in 2005 and almost 2% of the world GDP. Japan, for it turns, has long been the country with the largest current account surplus in the world (until 2006 when China is expected to have the largest surplus).

[25] In terms of the model's parameters, this would be represented by time varying risk aversion, intertemporal elasticity of substitution, and/or a subjective discount rate.

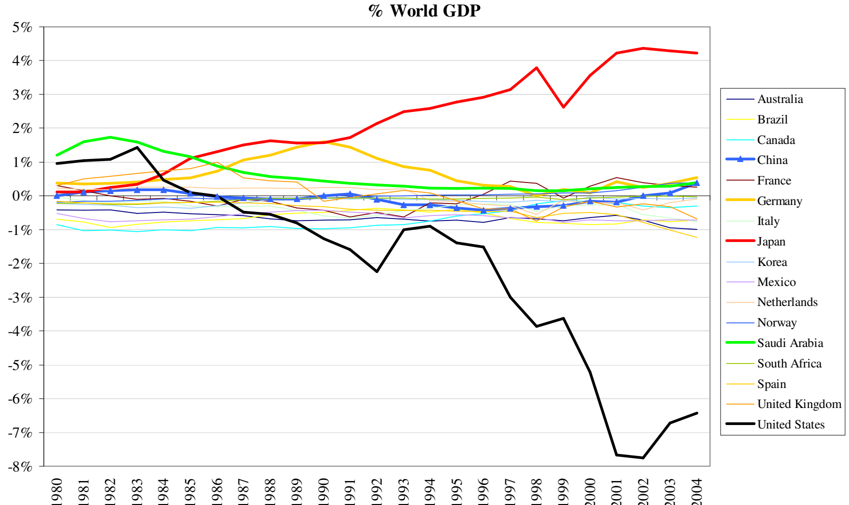
[26] See Eichengreen (2006) and Bernanke (2005) among many others.

Figure 4
Current Account Balance
as % of World GDP



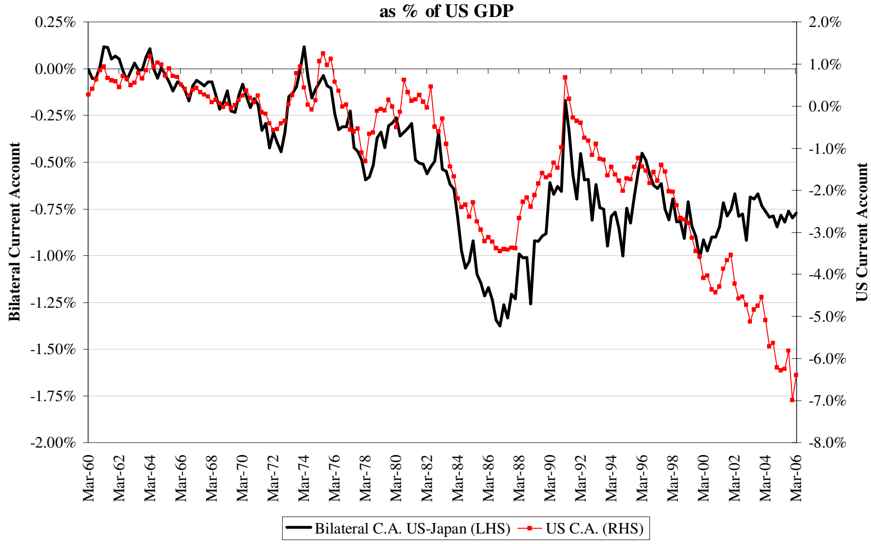
The importance of the US and Japan is even bigger when their external wealth is considered. As pointed out by Lane and Milesi-Ferretti (2005), Japan is by far the largest net creditor in international investment position and the US is by far the largest net debtor international investment position. These figures are presented on Figure 5.

Figure 5
Net External Wealth
% World GDP



Further, the total US current account and the US-Japan bilateral current account have similar dynamics. Thus, a theory for the US-Japan bilateral current account could also be relevant to the total US current account. This is highlighted by the Figure 6, where we compare the total US current account and the bilateral current account US-Japan. An explanation for the dynamics of the US-Japan bilateral current account could be an explanation for the total US current account dynamics.

Figure 6
US Current Account
as % of US GDP



On the discussion of global imbalances many competing explanations have been raised: the "global savings glut" as in Bernanke (2006), the low US savings/Twin Deficits as in Roubini (2005), Roubini and Sester (2004, 2005), and better investment opportunities as in Cooper (2004) and Backus and Lambert (2005). Our model can be understood as a formalization of this hypothesis and our estimated results as an empirical test of the relevance of this explanation to US current account imbalances. Our results indicate that better investment opportunities seem to be an important explanation for the phenomenon.

4.1. Estimating and Solving the Model

Results presented here take the perspective of an American investor as the representative agent who makes a portfolio choice over different assets from the US and from Japan.[27] In the Appendix, we report the results for a Japanese investor that faces a symmetric problem, but results are very similar.

We included as many asset classes as possible on the model. Data was available for stock market returns, short-term government bond rates, long-term government bond rates, and private firms' profits in a quarterly basis for both countries since 1960. The source for short-run interest rates and stock market returns data is Bloomberg. ROE is constructed as total operational profits/capital (net worth). The sources are the financial statements from Ministry of Finance for Japan and the US Flow of Funds Accounts calculated by the Federal Reserve for the US. Long-term interest rates data comes from Global Financial Database. We have also included in the VAR variables known to predict future returns such as price-to-earnings ratio and the US nominal short-term yield, which comes from Global Financial Database. All variables are measured in natural units so that standard deviations are per quarter.

The estimated VAR imposed the restriction that the unconditional means of the variables implied by the coefficient estimates are equal to their full-sample arithmetic

[27] Ideally, we would want to model the decisions of both and US investor and a Japanese investor in a single framework. However, at this stage, we are still working on how to aggregate these two types of investors.

counterparts[28]. The results[29] are presented in Table 7. The first row corresponds to the US Real T-Bill rate, the second row to the US ROE equation, and so forth. The coefficients of the estimation are taken as given and known by investors.

Table 7
VAR Estimation Results

		rtb_t	$usxroe_t$	$jpxsr_t$	$usxlr_t$	$usxs_t$	$jpxlr_t$	$jpxs_t$	$jpxroe_t$	$usxsr_t$	$usep_t$	$jpep_t$	R-Squared	F-Statisti	P-Value
US Real T-Bill	rtb_{t+1}	0.10	0.33	-0.19	0.23	0.00	0.17	0.00	0.15	1.32	0.02	0.00	0.89	286.52	0.00
		0.67	4.06	-3.94	3.27	0.73	4.33	0.64	3.06	6.19	1.07	0.20			
US ROE	$usxroe_{t+1}$	0.17	0.82	0.31	-0.09	0.00	-0.18	0.00	-0.29	-0.47	0.00	0.00	0.94	530.96	0.00
		1.34	8.16	5.18	-1.03	-0.97	-3.83	0.03	-4.68	-2.08	0.17	0.65			
JP Short-term Interest Rate	$jpxsr_{t+1}$	1.01	-0.28	0.09	-1.22	0.00	0.55	-0.02	0.08	-1.66	-0.09	0.01	0.28	13.519	0.00
		1.56	-0.69	0.28	-3.41	-0.14	2.25	-1.87	0.23	-1.69	-0.88	0.67			
US Long-term Bond	$usxlr_{t+1}$	0.25	-0.25	0.24	0.71	0.00	-0.17	0.00	-0.23	-0.70	-0.01	0.00	0.74	96.239	0.00
		2.60	-3.06	4.97	9.83	-0.71	-4.45	0.54	-4.48	-4.17	-1.01	0.95			
US Stock	$usxs_{t+1}$	8.29	-1.62	3.01	-1.31	0.04	-4.80	-0.07	-2.63	-17.94	-2.60	0.07	0.35	18.83	0.00
		2.79	-0.72	1.66	-0.62	0.53	-2.74	-1.02	-1.40	-3.59	-4.57	0.63			
JP Long-term Bond	$jpxlr_{t+1}$	0.29	-0.27	0.20	-0.25	0.00	0.83	0.00	-0.17	-0.69	0.00	0.00	0.93	441.39	0.00
		2.37	-2.60	3.54	-3.09	-0.29	16.47	0.01	-2.91	-3.16	-0.26	0.34			
JP Stock	$jpxs_{t+1}$	1.43	-1.81	-5.28	-2.62	0.39	3.12	0.22	4.13	-2.39	-0.36	0.09	0.21	9.0279	0.00
		0.37	-0.59	-2.07	-1.10	4.81	1.07	2.40	1.54	-0.36	-0.48	0.51			
JP ROE	$jpxroe_{t+1}$	1.01	-0.21	-0.77	-1.11	0.00	0.44	-0.02	0.90	-1.73	-0.11	0.01	0.22	9.6042	0.00
		1.57	-0.53	-2.46	-3.14	0.11	1.80	-1.49	2.71	-1.78	-1.08	0.40			
US Nominal Short-term yiel	$usxsr_{t+1}$	-0.17	0.12	-0.29	0.10	0.00	0.17	0.00	0.27	1.37	-0.01	0.00	0.92	389.39	0.00
		-1.30	1.19	-5.32	1.29	1.06	4.46	0.11	4.75	6.02	-0.40	-0.05			
US E/P Ratio	$usep_{t+1}$	0.59	-0.35	0.33	-0.22	0.00	-0.46	0.00	-0.28	-1.55	0.82	0.00	0.92	417.53	0.00
		2.09	-1.93	3.68	-1.32	0.17	-3.64	-0.52	-2.68	-4.27	19.54	-0.53			
JP E/P Ratio	$jpep_{t+1}$	0.97	-0.78	0.65	-0.34	0.01	-0.85	0.00	-0.92	-2.90	-0.17	0.93	0.94	514.08	0.00
		1.48	-2.00	2.28	-0.92	0.55	-3.13	0.19	-3.00	-3.16	-1.40	27.43			

Cross-Correlation of Resilduals

	rtb	usxroe	jpxsr	usxlr	usxs	jpxlr	jpxs	jpxroe	usxsr	usep	jpep
rtb	1.00	-0.70	-0.07	-0.69	0.01	-0.73	0.08	-0.07	0.76	0.06	-0.01
usxroe		1.00	0.07	0.79	0.14	0.81	-0.11	0.08	-0.96	0.10	0.05
jpxsr			1.00	0.09	-0.08	0.14	0.55	0.99	-0.04	-0.05	-0.12
usxlr				1.00	0.10	0.77	-0.04	0.09	-0.78	0.07	0.07
usxs					1.00	0.16	-0.03	-0.08	-0.09	0.48	0.24
jpxlr						1.00	-0.01	0.13	-0.85	0.09	0.04
jpxs							1.00	0.55	0.14	-0.06	-0.02
jpxroe								1.00	-0.05	-0.05	-0.11
usxsr									1.00	-0.08	-0.03
usep										1.00	0.21
jpep											1.00

The estimated VAR reports standard effects identified in the finance literature. US real short-term rate is significantly explained by the short-term nominal yield with a positive coefficient. US long-term bond rate is significantly explained by US real and nominal short-term rates. While the real rate is positively related to the long-term rate, the nominal rate is negatively related. US stock returns are negatively related to both nominal short-term yields and price-to-earnings ratio. The results for the Japanese rates are also similar. All equations are significant at the usual level as it can be seen from very high values of F-statistics. However, not all of them have high R-squared. Predicting excess stock returns is difficult. This can be seen from the R-squared from these equations. They are the lowest ones, between 0.21 and 0.28. Variables that help predict returns have a process close to an AR(1), with very high auto-regressive coefficients.

The model uses the information on expected returns. Summary statistics of the expected returns implied by the estimated VAR are reported on Table 8. Japanese assets have a higher variance than correspondent US assets. This could is probably reflecting an the exchange rate risk carried by these assets, since we are dealing with the US investor point of view. The time series of these expected returns are presented in Figure 7. As it can be seen in Figure 8, in the beginning of the sample from the 60s to mid 70s Japanese expected returns were higher. This figure was reversed later as

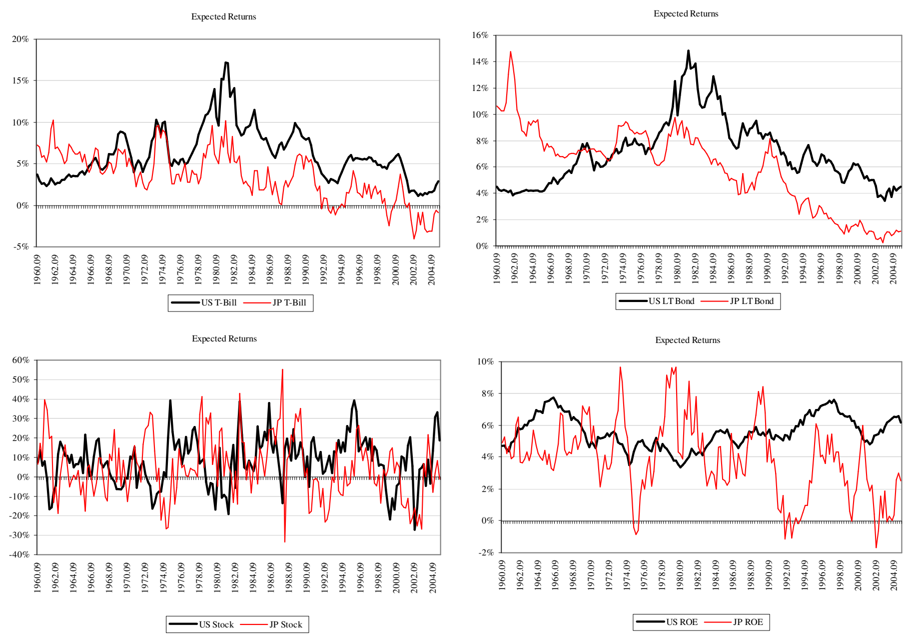
[28] No Bayesian priors were used in this estimation.
[29] Though finite sample bias might have some effect on the reported coefficients, bias corrections are complex in a multivariate system, so no corrections were attempted.

from the mid-80s onwards US expected returns have been consistently higher than Japanese expected returns.

Table 8
Summary Statistics of Expected Returns

Variables	Mean	Std. Dev.	AR(1) Coeff.	Std Error
US Real Short-Term Bond	5.04%	0.026	0.952	0.018
US Long-term Bond	7.06%	0.025	0.976	0.013
US Stock	7.76%	0.124	0.641	0.058
US ROE	5.57%	0.010	0.964	0.020
JP Short-Term Bond	3.48%	0.030	0.892	0.033
JP Long-term Bond	5.94%	0.031	0.988	0.012
JP Stock	4.62%	0.157	0.483	0.055
JP ROE	3.86%	0.024	0.825	0.038

Figure 7



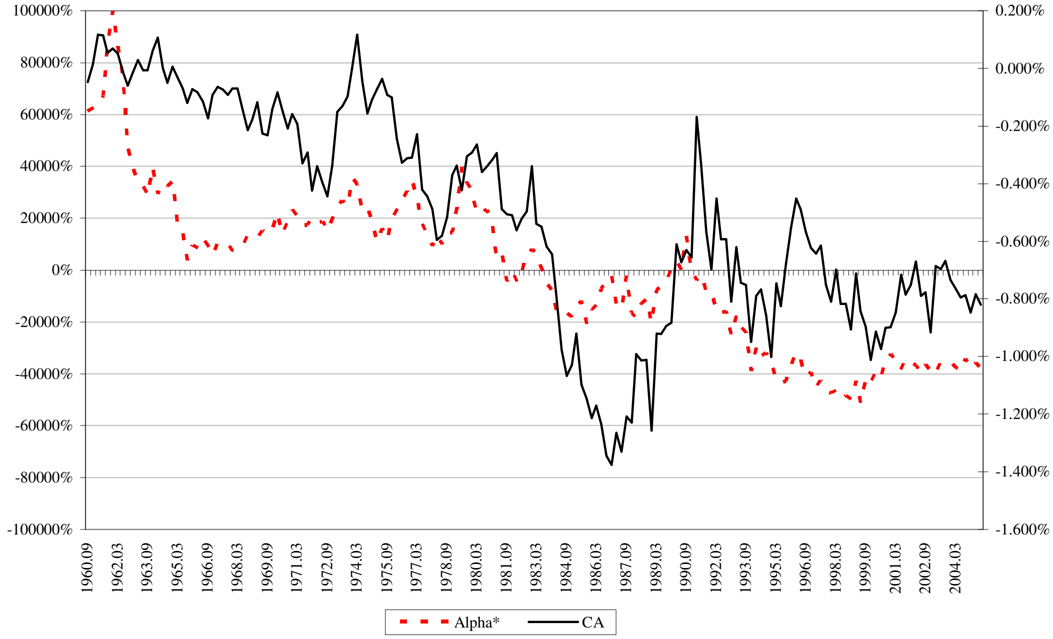
4.2. Explaining the US-Japan bilateral current account

Using the estimated coefficients, the model is calibrated for different risk aversion parameters (γ) and different intertemporal elasticity of substitution parameters (ψ). The reported results assume that $\psi = 1$ and $\delta = 0.92$ in annual terms. The results are robust to different parameter values. Now, it is possible to understand how well this dynamic portfolio model can explain the current account dynamics. We calculate the fraction of the optimal portfolio allocated abroad (if we consider an American investor) for each data point (quarter). Our main variable of interest is, therefore, the proportion of US wealth that should be invested in Japan, which is the sum of the US the optimal proportion of wealth to be invested in each Japanese asset class. Formally,

$$\alpha_t^* = \alpha_{stocks,t}^* + \alpha_{ROE,t}^* + \alpha_{Bonds,t}^*$$

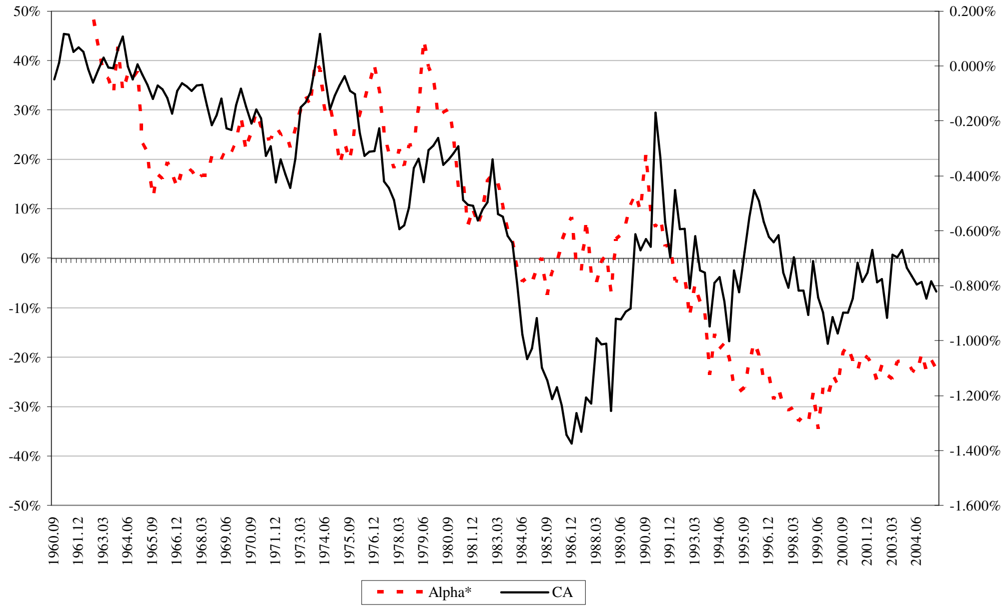
Figures 8 and 9 plot the time series of the optimal fraction of wealth to be allocated abroad and the bilateral US-Japan current account. Remember that these two variables are related through the basic current account identity[30]: $CA_{t+1} = \alpha^*_{t+1}W_{t+1} - \alpha^*_t W_t$. The theoretical identity tells us that when the optimal portfolio share allocated to Japanese assets fall, everything else constant, the bilateral current account should worsens as well. The fitting of the graphs is good, especially if one consider that only information on asset returns were used. In other words, the model did not use any information on the current account dynamics. The results are robust to the inclusion/exclusion of assets and to different parameter values.

Figure 8
Optimal Portfolio Share Abroad and Bilateral Current Account:
US Representative Agent and Gamma = 2



[30] Actually, in general equilibrium, the true identity would be $CA_{t+1} = [\alpha^*_{t+1}W_{t+1} - \alpha^*_t W_t] - [\alpha^{JP}_{t+1}W^{JP}_{t+1} - \alpha^{JP}_t W^{JP}_t]$ where α^{JP}_t is the US portfolio share of a Japanese investor. Thus, the figures are actually addressing half of the story. However, this is not a problem, for the contrary: in the appendix we show the similar figures for a Japanese investor and the behavior of Japanese agents is very similar to its US counterpart. This implies that the results we find here are reinforced by the foreigners' side of the story. Since we are still working on how to aggregate these two types of agents, we preferred to focus on the domestic-determined side of the current account for now without loss of generality.

Figure 9
Optimal Portfolio Share Abroad and Bilateral Current Account:
US Representative Agent and Gamma = 2000



The main mechanism behind the re-allocation of portfolio shares in our model is the expected changes in risk premia over the different assets. Suppose the expected return of US bonds increases while all other expected returns remain unchanged. This represents an improvement in the investment opportunity set. According to the dynamics of the model, an investor should then increase its portfolio share on US bonds. If this investor is a Japanese investor, this implies that the US bilateral current account would worsen. If a US investor is considered, this implies that her shares on Japanese assets should fall, and hereby, improving the US bilateral current account as well. By putting together this mechanism and the data on expected returns plotted on Figure 7, it is clear that investors have been increasing their share of US assets, the reason being improvements in US investment opportunities.[31] These changes seem to be an important part of the explanation for the observed shifts towards US assets of country portfolios in the last 15 years. Thus, this is strong evidence that the mechanism analyzed is indeed an important one.

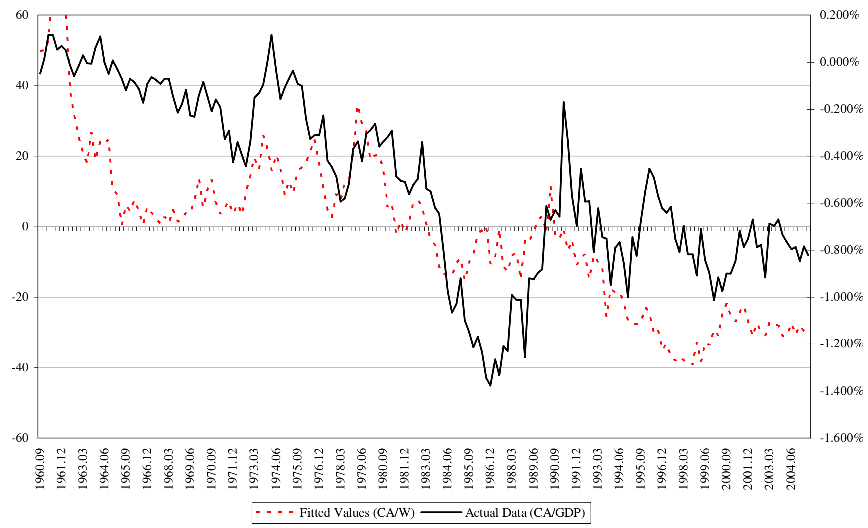
As it can be seen in the graphs, variations in the optimal portfolio do not change considerably when gamma changes. But it is worth noting that the absolute value of the proportion of wealth allocated abroad does vary significantly with gamma. When small gammas are used, the fractions are extremely high. Reasonable values of the proportions of the portfolio invested in each asset class are obtained only when very large values are considered (gammas as high as 2000). This happens because gamma is capturing two effects in place. It captures the well-known equity premium puzzle, in which high values of gamma are needed to obtain reasonable values of the excess return of equity over bond interest rates. It also captures a home bias effect (or the border effect) that it is well documented in the international finance literature. In order to induce agents to hold small shares of Japanese assets, characterizing home bias, we need to make them extremely risk averse. Even though the procedure does miss the average level of the current account for small values of gamma, it still captures well

[31] In the Appendix, we show the composition of portfolio shares over different assets over time.

the dynamics of the series.[32] Thus, it does achieve our main purpose of explaining the movements on the bilateral current account. Including financial constraints such as borrowing constraints in the model could minimize it, but by doing so, the model would become extremely hard to solve. This feature can let us conclude that, in order to explain the observed current account dynamics, the representation of asset returns in a VAR form is more important than the use of the Epstein-Zin utility function.

We can take a step forward and try to calculate the current account predicted by our model. Again, we do so by looking at the behavior of domestic agents only. Given the evidence we have at this point on the behavior of a Japanese agent, introducing them to the picture would only reinforce our results. Figure 10 plots the results for gamma equals to 20, a reasonable value when the savings decision does matter to the analysis. We scaled our predicted measure of the current account by wealth, the stock variable of our model. The actual data is scaled by GDP instead, since there is no US wealth data available at a quarterly basis. As before, the fit is impressive.

Figure 10
Bilateral Current Account
Gamma = 20



More formally, an econometric analysis is able to confirm the intuition from the figures. A regression of the estimated optimal portfolio share invested abroad on the US-Japan bilateral current account always shows a positive and significant coefficient. The results are reported on Table 9. The lowest R-squared obtained from our model is 0.54. This is in sharp contrast with the very low R-squared obtained in other time-series regressions for constant portfolio models, presented in Table 5 of Chapter 3. This implies that variations in investment opportunities could be responsible for at least 54% of the pattern observed in US-Japan bilateral current account.

However, the current account series is highly persistent as it can be seen by the low Durbin-Watson statistics. In order to take that into account, a specification including the lag of current account is also performed. The results remain qualitatively unchanged as our measure of optimal portfolio abroad is still highly

[32] This is not a problem inherent to our model. The original model by Campbell, Chan and Viceira (2003) also underwent similar difficulties.

significant and positive. As expected, in this specification, the R-squared are much higher, around 0.92.

Table 9
Dynamic Portfolio Model and Bilateral Current Account
Dependent Variable: US-Japan Bilateral Current Account

	Gamma = 1		Gamma = 2000	
Constant	-0.005	-0.001	-0.006	-0.001
	[26.805]***	[3.676]***	[30.074]***	[3.717]***
Alpha*	0.000	0.000	0.011	0.001
	[19.456]***	[3.127]***	[19.319]***	[3.137]***
BCA(t-1)		0.893		0.895
		[26.917]***		[27.538]***
No. of Obs.	180	179	180	179
R-squared	0.559	0.924	0.544	0.924
D-W Stats	0.193	2.288	0.187	2.292

Robust t statistics in brackets
* significant at 10%; ** significant at 5%; *** significant at 1%

We also implemented a co-integration analysis. The results are on Table 10. There is some evidence of a co-integrated relationship between the bilateral current account and the optimal portfolio share allocated abroad. The error correction model indicates that the bilateral current account reacts to a deviation of their long run relationship. In other words, an increase in the optimal portfolio share leads to a significant increase in the bilateral current account.

Table 10
Dynamic Portfolio Model and Bilateral Current Account
Dependent Variable: US-Japan Bilateral Current Account

Cointegration Equations

	Gamma = 1	Gamma = 2000
Constant	0.0049831	0.0063526
Alpha* (t-1)	-3.47E-06	-0.0143503
	[-3.85496] ***	[-3.83319] ***
BCA (t-1)	1.000	1.000

Vector Error Correction Estimates

	Gamma = 1		Gamma = 2000	
	D(CA)	D(Alpha*)	D(CA)	D(Alpha*)
Coint. Eq.	**-0.07**	5582.16	**-0.07**	1.40
	[-2.34]**	[0.89]	**[-2.35]****	[0.91]
D(Alpha* (t-1))	0.00	-0.03	0.00	-0.03
	[-0.54]	[-0.39]	[-0.54]	[-0.37]
D(BCA (t-1))	-0.17	-5225.64	-0.17	-1.79
	[-2.24]**	[-0.34]	[-2.27]**	[-0.45]
C	0.00	-19.98	0.00	0.00
	[-0.74]	[-1.24]	[-0.74]	[-1.22]
No. of Obs.	178	178	178	178
R-squared	0.07	0.01	0.07	0.01
F-statistic	4.48	0.39	4.49	0.43

Robust t statistics in brackets
* significant at 10%; ** significant at 5%; *** significant at 1%

We had already provided evidence that country portfolios varies over time and, thus, drives the current account accordingly. The results of this section lead us to the conclusion that varying investment opportunities seems to be a very important mechanism behind country portfolios and current account changes. The model proposed have better empirical results than constant portfolio models. The R-squared on the time-series regressions are higher and none of the model's hypothesis was violated in the estimation. To our knowledge, this is the first dynamic portfolio model for the current account in the international economic literature. One important drawback of our approach is that, mostly due to data restriction, we were only able to examine US-Japan bilateral relationship. However, as already argued, these two countries seem to be the most important ones to explain global patterns.

5. Conclusion

The current account balance can be understood as an asset allocation decision: how much of the national savings should be invested abroad. Surprisingly though, theoretical models did not incorporate this important asset allocation feature of the problem in a clear way until very recently. However, even then, the optimal portfolio was considered constant over time. Moreover, the constant portfolio models could explain on average only 18% of the current account movements. In the present paper, we argue that country portfolios vary over time and this should not be ignored.

We propose a dynamic portfolio model for the current account with time varying investment opportunities and estimated it for US and Japan. In our empirical analysis, we compared the time series of the optimal fraction of wealth allocated abroad with the corresponding bilateral current account. Econometric tests provided some evidence of a positive and significant relation between these two series. This model was able to explain at least 54% of the US-Japan bilateral current account movements, a much higher figure than in the constant portfolio models. Thus, country portfolios seem to vary over time and our results indicate that changes in investment opportunities are a very important mechanism behind it.

In the context of the current discussion of "Global Imbalances", many competing explanations, especially for the rising US current account deficits, were developed. One of them relies on the argument of better investment opportunities in the US compared to other G7 countries as in Cooper (2004), Clarida (2005), and Backus and Lambert (2005). In this paper, we formalize this argument and introduce it into an intertemporal model for the current account. As shown by Lane and Milesi-Ferretti (2005), Japan is by far the largest net creditor in international investment position and the US, the largest net debtor in international investment position. Our results indicate that a significant part (but not the total) of the US current account imbalances can be explained by better investment opportunities.

Extensions to the model are related to the development of an integrated framework to include both US and Japanese agents. Other developments to the model could be related to financial constraints, such as short-sale constraints or limited ability to invest abroad. On the empirical side, an obvious extension is the inclusion of other countries in the analysis, such as European and Asian countries. As highlighted by Gourichas and Rey (2006), in our model, we can distinguish valuation effects on current holdings from genuine changes in portfolios. Thus, the empirical relevance of valuation effects on current holdings and the consequent need to rebalance also deserves some attention.

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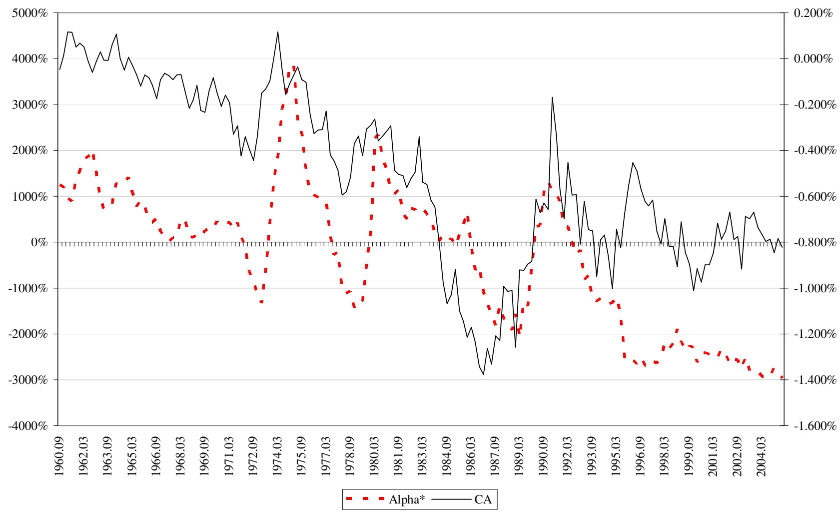
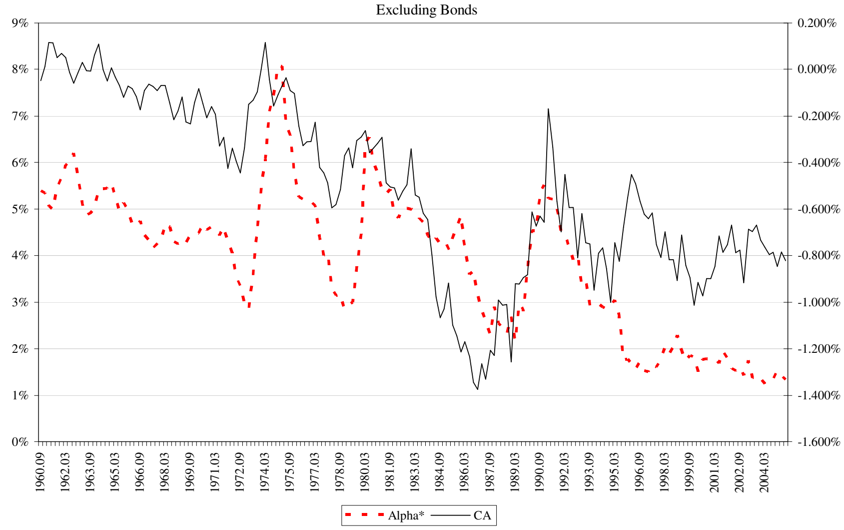
Appendix

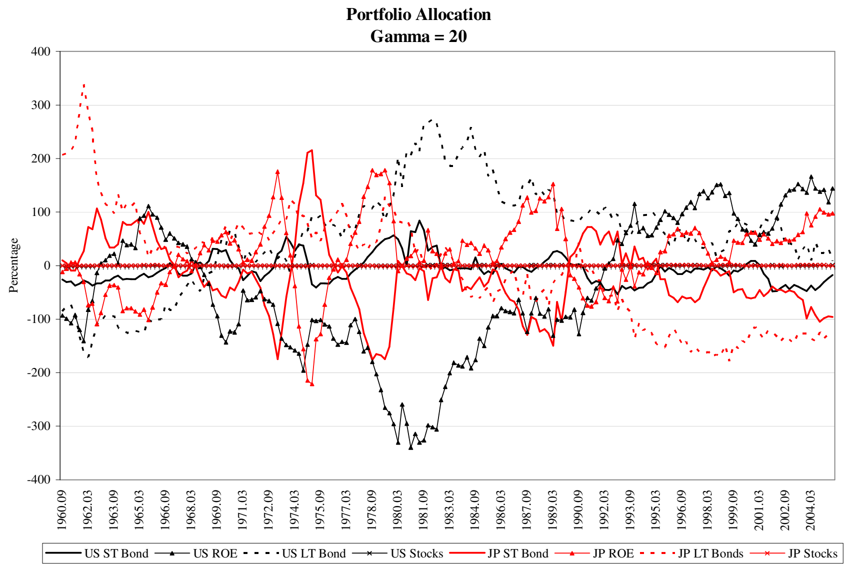
A. Japanese Agents’ Problem

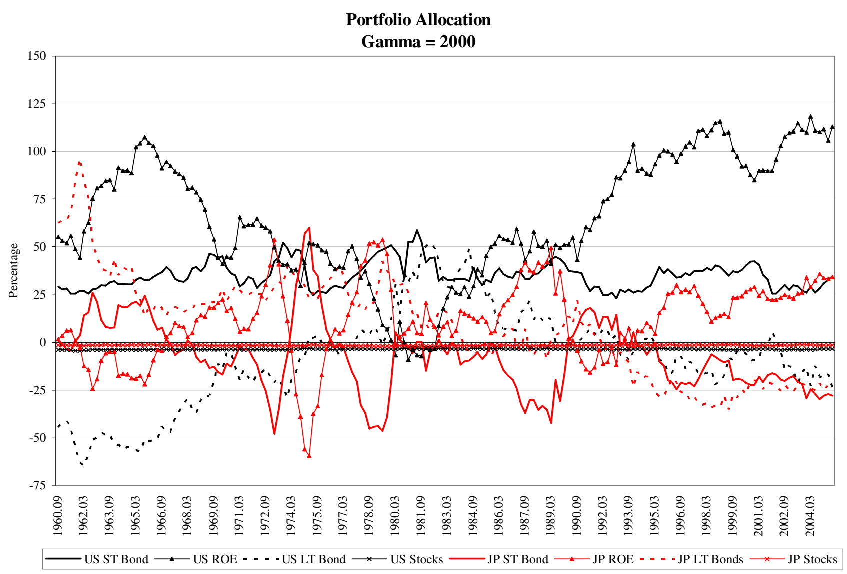


Optimal Portfolio Share Abroad and Bilateral Current Account:
Japanese Representative Agent and Gamma = 2000
Excluding Bonds



B. Optimal Portfolio Composition





Cousin Risks: The Extent and the Causes of Positive Correlation between Country and Currency Risks*

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Abstract

When country and currency risk premiums are positively correlated, a negative international liquidity shock harms twice the small open economy, thereby substantially increasing interest rates. This harmful positive correlation between country and currency risk premiums observed in some countries is called cousin risks. We, first, identify the extent of this phenomenon by separating a sample of countries into two groups: the one where the positive correlation is observed and the one where it is not. Based on this taxonomy, we investigate the determinants of the cousin risks. Results indicate that currency mismatch and low financial deepening are strongly associated with the phenomenon.

Key words: Country Risk, Currency Risk, Financial Crisis, Interest Rate, Cousin Risks

JEL classification: E43, G15, F34

1 Introduction

In times of reversal of capital flows and worldwide economic slowdown, as in 2001 and 2002, a few emerging markets are burdened with higher real interest rates precisely when growth is faltering. This combination of bad outcomes constitutes the opposite of the smoothing effect that financial markets are expected to provide. However, the impact of the reversal of capital flows is felt differently across emerging

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markets, as some countries are more vulnerable than others. In order to overcome these fragilities, it is imperative to identify their sources[1].

The covered interest rate parity (CIP) condition can be used to decompose the domestic interest rate into three components: the international interest rate, the forward premium, and a residual that proxies for the sovereign credit risk premium (the so called country risk). The forward premium—measured by the difference of the log of the forward exchange rate and the log of the spot exchange rate—encompasses both the expected depreciation, and the currency risk premium. The joint behavior of country risk and forward premium can be used to analyze the effect of shocks to both the supply of and the demand for international capital flows[2]. Under this framework, vulnerability to external shocks is identifiable through the high level and volatility of both premiums.

An additional vulnerability source occurs up when a country presents positive correlation between country risk and currency risk. That is because, given the CIP, shocks on those two components would occur at the same time and in the same direction, magnifying the necessary interest rate reaction to avoid capital flight. Contrasting with the myriad of papers that aim at understanding how each of these two risks behaves separately, the ones that focus on their co-movement, as the present work does, are scarce.

Powell and Sturzenegger (2000) analyzes if there is a causality relation between currency risk and country risk in a small sample and conclude that the patterns are quite diverse. Garcia and Didier (2003) identified a large and positive correlation between the two risks in Brazilian data. The authors credited this result to the fact that those risks share a common generating factor. Due to the likely existence of a common root for the two risks, the authors named them cousin risks. Deepening that line of research, our paper has two goals. The first one relates to the analysis of the correlation pattern of those two risks among a sample of 25 countries[3], while the second one aims at finding the factors that are behind their common root. In short, we will first investigate how widespread the cousin risk phenomenon is. Having identified its prevalence, we will go on to examine the possible causes of the positive correlation between country and currency risk premiums.

This paper has five sections. Section 2 puts the term cousin risk in context. Section 3 investigates how widespread the cousin risks phenomenon is. Having identified the extent of the cousin risks phenomenon, Section 4 studies the determinants of the cousin risks. Section 5 concludes.

[1] Recent models have tried to identify what makes some economies "financially vulnerable" while others remain "financially robust", e.g. Krugman (1999), Aghion, Bacchetta and Banerjee (1999, 2001 and 2004), Caballero and Krishnamurthy (2000 and 2001a to c), Caballero and Panagea (2003), Calvo and Reinhart (2002), Christiano, Gust and Roldos (2002).

[2] Henceforward, we shall drop the word premium and refer simply to country risk and currency risk.

[3] Such an empirical study only recently became possible, since it presupposes the existence of forward exchange rate markets in different currencies which is the binding restriction to construct our sample, that so far has 25 countries.

2 Cousin Risks

2.1 Covered Interest Rate Parity Condition

Under free international capital flows, the covered interest parity (CIP) condition[4] states that the domestic interest rate may be broken into two components: the international interest rate and the forward premium:

$$1 + i_t = (1 + i_t^*)(\frac{f_t}{s_t}) \Rightarrow i_t \cong i^* + (FwdPremium_t) \tag{1}$$

Where, i_t is the internal interest rate of a domestic bond denominated in the domestic currency, from t to t+1; i_t^* is the risk free international interest rate from t to $t+1$; f_t is the forward exchange rate traded in t to be delivered in $t+1$; s_t is the spot exchange rate in time t.

When there exists default risk, we need to include in the parity condition above a credit risk premium charged by the investors. Denoting θ_t the country risk (sovereign default risk premium), the parity condition becomes:

$$1 + i_t = (1 + i_t^*)(\frac{f_t}{s_t})(1 + \theta_t) \Rightarrow i_t \cong i^* + (FwdPremium_t) + (CountryRisk_t) \tag{2}$$

Studies of the forward premium have traditionally used interest differential between countries to proxy for the forward premium. This approach assumes absence of country risk, which generaly does **not** hold for the emerging markets. Recently, the development of derivative markets for emerging market currencies has rendered possible the direct calculation of forward premiums on a daily basis.

As Fama (1984), Bansal and Dahlquist (2000) and several other authors show, it is a stylized fact that forward exchange rates are biased estimators of the actual spot exchange rate in the future, a puzzle known as the forward premium puzzle. The literature considers many possible explanations for the forward premium puzzle: existence of a risk premium, market inefficiency, lack of rational behavior, learning, peso problem, and others. Assuming that the first risk premium explanation is true, the forward premium is equal to expected depreciation plus a currency risk premium, which, in turn, is a result of exchange rate uncertainty.

The measurement of this unobservable risk premium is not a trivial task, since all that forward market quotes provide is the sum of currency risk with expected depreciation, i.e., the forward premium. To separate the two components of the forward premium, one must rely on extra information, as surveys of market expectations or statistical frameworks[5]. Since both ways could introduce extraneous noise in our procedure, we choose to concentrate on the forward premium as a whole, i.e., on the expectation of depreciation and the risk premium relating to its uncertainty. Thus, henceforth "currency risk" and "forward premium" will be used interchangeably. In our sample we calculate the currency risk (i.e., the

[4] Frankel (1991), Domowitz et al (1998) and Garcia and Didier (2003) are some papers that use this framework.

[5] Wolff (1987, 2001) and Garcia and Olivares (2001).

forward premium) using data[6] on 1 year forward exchange rate and on spot exchange rate as follows:

$$ForwardPremium_{1year,t} = \frac{(\text{forward rate}_{1year,t} - \text{spot rate}_t)}{\text{spot rate}_t} \tag{3}$$

The country risk premium is relevant when agents perceive a possibility of default. We calculate it by two procedures: (1) EMBI+ or EMBI GLOBAL spread and (2) covered interest parity differential (CID). CID is the interest rate deviation vis-à-vis the value predicted by the non-arbitrage condition stated by the CIP on de absence of credit risk. We calculate the CID using the 1 year Swap rate in local currency as i_t and the 1 year US Treasury rate in dollars as i_t^*:

$$CID_t = i_t - i_t^* - (ForwardPremium)_t \tag{4}$$

Alternatively, we could measure a country's credit risk through its issued bonds denominated in a foreign currency. Such a bond would not be subject to currency risk since it is denominated in a foreign currency. It is, nevertheless, subject to the issuer's credit risk. Thus country risk would be equal to the implicit rate of this bond exceeding the international risk free interest rate of same duration, i.e.:

$$CountryRisk_t = i_t^{US} - i_t^* \tag{5}$$

Where, i_t^{US} is the interest rate of one of its issued bonds denominated in a foreign currency (usually the US dollar), from t to t+1; i_t^* is the international risk free interest rate from t to t+1. In our sample, we use directly the JPMorgan's EMBI+ and EMBI GLOBAL stripped spread[7].

The literature on the determinants of country risk is very large. Many papers resort directly to econometric modeling without an explicit model. The goal is to evaluate each variable's net effect over credit risk. Garcia and Didier (2003), Westphalen (2001), Kamin and von Kleist (1999) and Mauro, Sussman, and Yafeh (2000) are papers that follow this methodology. In all of the aforementioned papers, explanatory variables can be classified into three groups: 1) liquidity and solvency variables; 2) macroeconomic performance variables and; 3) global risk aversion variables. In group 1, the main variables affecting country risk are debt over GDP ratio, debt service over exports ratio, debt service over GDP ratio, and the level of international reserves. In group 2, the following variables stand out: GDP growth, inflation rate, and terms of trade. Lastly, the junk bond or high yield spread is largely used as a measure for global risk aversion.

2.2 Why would country and currency risks follow a similar trend?

The literature on the co-movement of the forward premium and the country risk premium is still very incipient. From a logical point of view, a strong correlation between any two series can only arise under

[6] Data sources are provided in Appendix 4.

[7] EMBI+ and EMBI global are indexes, constructed by JPMorgan, composed by the most liquid U.S. dollar-denominated bonds. EMBI's stripped spread is simply the difference between that index and a US Treasury rate of same duration.

one of two conditions: the first is the existence of a common generating factor, and the other possibility is the existence of a causality relation between the two series, i.e., movements in one series influencing the behavior of the other.

In regard to the first possibility, country risk and forward premium are analyzed in the literature and their respective individual determinants are widely researched. These would be the natural candidates of being a common factor, i.e., a factor that would have generated both series. Nevertheless the literature argues that each variable has different determinants. The main determinants of country risk are solvency and liquidity variables (e.g. level of net indebtedness, fiscal deficits, and global risk aversion), while the main components of forward premium dynamics are related to the balance of payments uncertainties. In Section 4 we will formally test if the occurrence of the positive correlation phenomenon is associated with a high (or low) level of these variables.

The causality relation has received some support in the literature. Can forward premium shocks trigger off country risk shocks? Powell and Sturzenneger (2000) try to answer this question. Using an event-study methodology they analyze the causality effect of currency risk on country risk. Their result indicate that there are various patterns. A few countries present positive relation while others present negative or no relation at all. Positive causality was found in Argentina, Austria, Belgium, Brazil, Ecuador, Ireland and Mexico; negative, in Denmark, Portugal and Sweden. Our section 3 presents an analysis of forward premium and country risk joint behavior for a larger sample of countries but our framework does not allow us to infer causality.

There are theoretical arguments in favor of both positive and negative relation between the two risks. If an economy dollarizes, the abandonment of national currency means the abolition of seigniorage and, as a consequence, a possible worsening of the country's fiscal conditions, increasing its credit risk. Another negative effect could come from the smaller nominal flexibility of a dollarized economy, which would imply higher real response to shocks, causing GDP's volatility to increase. In turn, this real volatility could increase country risk.

Conversely, there are arguments that justify a reduction of the country risk due to the abolition of the domestic currency, such as the increase in financial efficiency, the elimination the possibility of suffering speculative attacks, and the end of the government's balance currency mismatch. Increase financial efficiency, eases government funding, which could lead to uncertainty reduction regarding fiscal solvency, ultimately reducing the country risk.

The most interesting argument however is the so called balance sheet effect, which states that the effect of the forward premium on the country risk is due to government balance sheet currency mismatches. This currency mismatch occurs when a significant part of government liabilities are denominated in a foreign currency while assets and future proceeds are denominated in local currency. Under these circumstances, domestic currency depreciation could affect government balance sheet, potentially leading the government to default on its debt. Broadening the exchange rate crisis model, Krugman (1999) presents a model in which balance currency mismatches in firms' balance sheets help to explain an exchange rate crisis. Neumayer and Nicolini (2000) presents, theoretical arguments regarding the relation between balance currency mismatches and country risk.

The 'balance sheet' argument is in line with Eichengreen, Hausmann, and Panizza's (2002) observation of the original sin phenomenon, which states that most of the countries cannot borrow internationally in their own currency. They say that only a few countries, referred to as major financial centers, do not face this problem: the USA, countries in the EURO zone, the United Kingdom, Japan, and Switzerland. According to them:

> ...while the major financial centers issued only 34 percent of the total debt outstanding in 1993-1998, debt denominated in their currencies amounted to 68 percent of total . ... Developing countries accounted for 10 percent of the debt but less than one per cent of currency denomination in 1993-1998 period. This, in a nutshell, is the problem of original sin.

Indeed, Hausmann (2002) claims that the composition of the net stock of debt could explain why, in spite of Latin American fiscal improvement efforts during the 90s, there were no significant improvements in country risk measures.

Despite the fact that many theories justify, by different arguments, correlation between currency risk and country risk, none of the papers reviewed here carried out an empirical investigation on the determinants of the positive correlation between the two risk premiums[8]. Such an analysis will be carried out in Section 4, but the initial objective, to which we turn now, is to identify the extent of the cousin risks phenomenon.

3 How widespread is the cousin risks phenomenon?

3.1 The data

We now investigate the extent of the cousin risks phenomenon, through an analysis of the country and currency risks' joint behavior in a sample of 25 countries: Australia, Argentina, Brazil, Canada, Chile, Colombia, Czech Republic, Great Britain, Indonesia, Japan, Mexico, New Zealand, Norway, Peru, Philippines, Poland, Russia, Singapore, South Africa, South Korea, Sweden, Switzerland, Thailand, Turkey, and Venezuela. Data frequency is daily and the time frame analyzed is January 1995 to January 2004, but it varies substantially across countries according to the data availability. For the thirteen

[8] Eichengreen, Hausmann, and Panizza (2002) estimated which factors could cause an exchange rate mismatch, but they do not estimate if this stylised fact is associated with the correlation between country risk and risk premium.

countries in our sample for which JPMorgan computes the EMBIs, we analyze the relation between embi spread and forward premium and for the rest of them, we analyze the relation between CID and the forward premium[9].

Whenever available, we prefer to work with EMBI spreads, since these are rates less affected by monetary policy interventions than the domestic interest rate (measured by the Swap rate). Also, EMBIs are calculated from the country's most liquid external bonds. Furthermore, if investors change their preferences during the period of analysis, JPMorgan adjusts the sample accordingly, thus EMBIs accurately depict investors' risk perception. Garcia and Valpassos (1998) and Garcia (2002) indicate that for Brazil, although the two measures of country risk are closely related, CID responds a little slower than the EMBI+ spread does.[10].

3.2 Results

The following graphs indicate how different the patterns of joint behavior can be, what confirms that the cousin risks phenomenon is not pervasive.

From the graphs we can infer that there is a strong positive correlation between country risk and currency risk in countries like Brazil and Mexico while in others, such as Colombia and South Korea, the cousin risks phenomenon does not seem to occur. Graphs 1a, 2a, 3a and 4a present country and currency risks time series as well as the rolling window of their correlation coefficient. In Brazil and Mexico, country risk and currency risk curves follow almost identical paths while in Colombia and South Korea they do not. Moreover, the graphic evidence from scatter diagrams 1b, 2b, 3b and 4b confirm our preliminary diagnostic: the positive linear pattern in Brazil and Mexico is remarkable. Even though this result stands out clearly from the graphs we shall carry out a formal statistical analysis.

[9] For 10 of the countries in our sample, we could calculate both measures of country risk: EMBI spread and CID. So for each of these 10 countries, we can have two correlations: the correlation between the FP and embi spread and the correlation between FP and CID. As expected the correlation of these two measures is highly positve: 0.6. These figures are presented in the Appendix 3.

[10] Garcia and Valpassos (1998) analyze the evolution of CID and the C-Bond spread in Brazil (C-Bond spread is similar to the Brazil's EMBI spread) during the controlled exchange rate regime. Undoubtedly, there is a close relationship between these variables and a large mismatch between them should cause other economic variables such as the exchange rate and international reserves to move. During the period analyzed, in the event of bad shocks the C-Bond Spread was the first to jump, and covered-interest-rate-parity differential moved later, as domestic interest rate were raised to avoid further foreign reserves losses. Therefore, the increase in the difference between the C-Bond spread and the covered-interest-rate-parity differential in Brazil had served as a very good coincidental, and sometimes leading, indicator of currency crisis. This paper does not extend the above study to a broader set of countries. The results in Garcia and Valpassos (1998) and Garcia (2002) indicate that CID responds more slowly than the EMBI spread does. So, EMBI spread is more reliable for capturing quick changes in investors' risk perception on a daily basis.

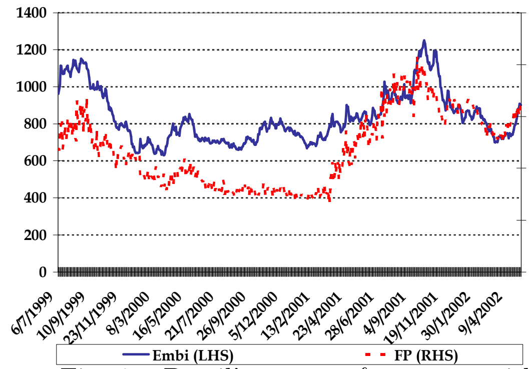

Fig. 1a: Brazil's country & currency risks

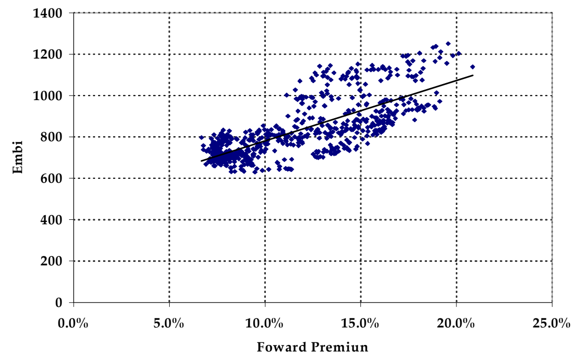

Fig. 1b: Brazil's risks scatter plot

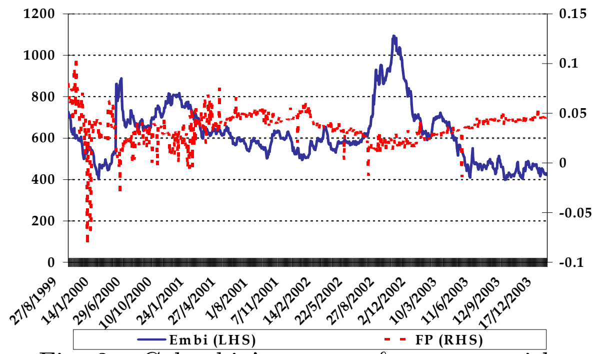

Fig. 2a: Colombia's country & currency risks

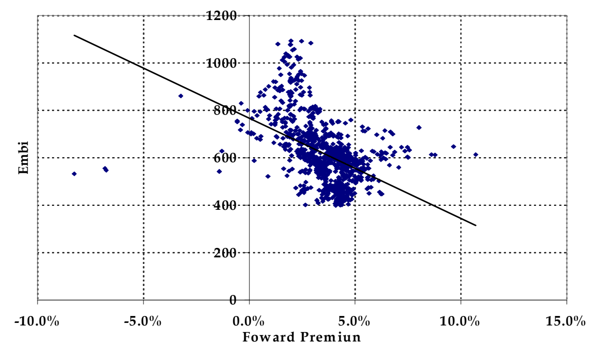

Fig 2b: Colombia's risks scatter plot

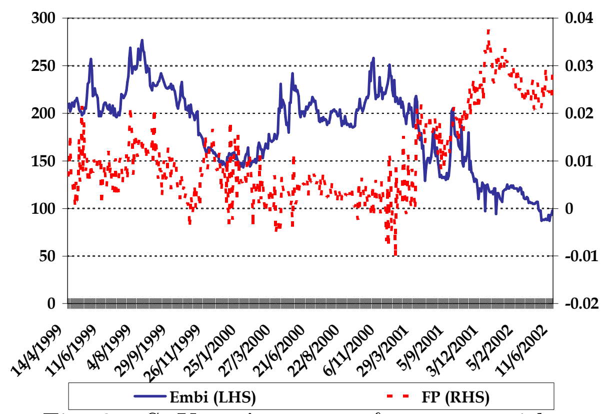

Fig. 3a: S. Korea's country & currency risks

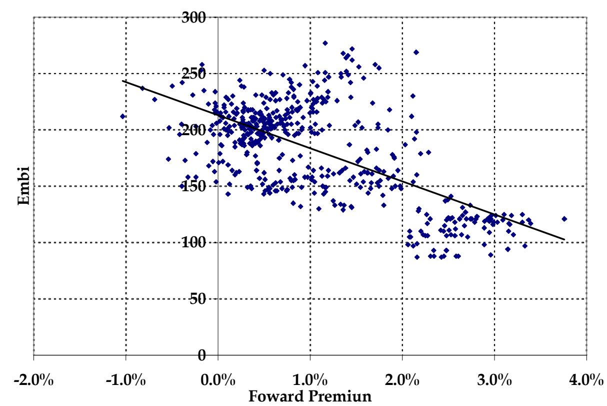

Fig. 3b: S. Korea's risks scatter plot

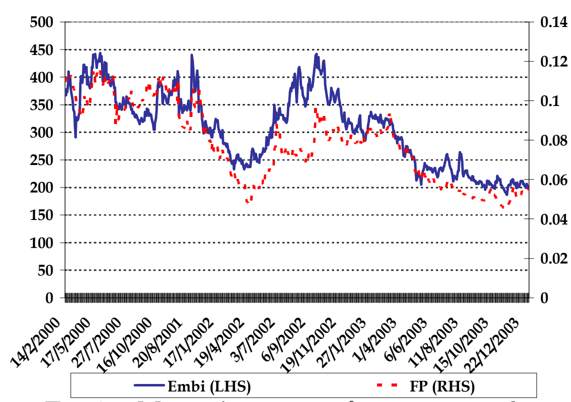

Fig 4a: Mexico's country & currency risks

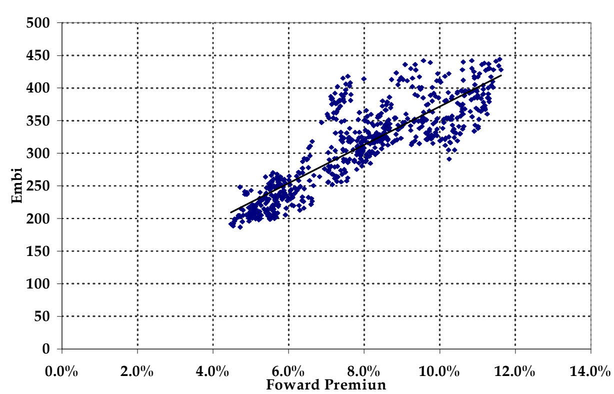

Fig 4b: Mexico's risks scatter plot

Table 1 presents the correlation coefficients, p-values and cointegration analysis results. A positive correlation is an indication of the presence of cousin risks phenomenon. Most of the series are non-stationary or close to non-stationary so the analysis needs extra care since the estimated coefficients can be spurious. We rely on two methods to identify the Cousin Risks phenomenon: non-parametric tests on the correlation coefficients (based on Monte Carlo simulation and bootstrap) and cointegration analysis.

Table 1 - Correlations, Cointegrations and p-values for H0 of Non-Positive Correlation (Monte Carlo and Boostrap)

	Data Sample		Correlation			Cointegration (at 1% significance level)?	Cousin Riks?
	Period Analyzed	Observations		p-value (Monte Carlo)	p-value (Bootsrap)		
1 S. Africa*	Feb/95 - Dec/03	600	**0,0585**	0,4487	0,4380	No	No
2 Australia	Jan/95 - Dec/03	2.191	**-0,8195**	0,9999	0,9999	-	No
3 **Argentina***	Jul/00 - Dec/01	334	**0,9267**	0,0007	0,0010	Yes - Positive	Yes
4 **Brazil***	Jul/99 - May/02	651	**0,7400**	0,0003	0,0000	Yes - Positive	Yes
5 Canada	Jan/95 - Dec/03	2.190	**-0,5684**	0,9999	0,9999	-	No
6 **Chile***	Jul/00 - Dec/03	816	**0,4572**	0,0000	0,0000	-	?
7 Colombia*	Aug/99 - Oct/02	984	**-0,4663**	0,9999	0,9999	-	No
8 Czech Rep	May/97 - Dec/03	1.328	**-0,8444**	0,9980	0,9999	No	No
9 Indonesia	Sep/96 - Mar/01	1.027	**-0,7738**	0,9999	0,9999	-	No
10 Japan	May/95 - Dec/03	2.155	**-0,7611**	0,9999	0,9999	-	No
11 S. Korea*	Mar/99 - Dec/03	491	**-0,6351**	0,9999	0,9999	-	No
12 **Mexico***	Nov/97 - Oct/02	682	**0,8609**	0,0000	0,0000	Yes - Positive	Yes
13 New Zealand	Jan/95 - Dec/03	2.205	**-0,7348**	0,9999	0,9999	-	No
14 Norway	Dec/95 - Dec/03	1.950	**-0,4556**	0,9999	0,9999	-	No
15 **Peru***	Jul/00 - Oct/02	843	**0,7162**	0,0000	0,0000	-	Yes
16 **Phillipines***	Mar/99 - Out/02	1.134	**0,7521**	0,0000	0,0000	-	Yes
17 Poland*	Jun/00 - Dec/03	1.070	**0,2535**	0,1285	0,1860	No	No
18 **Russia***	Dec/99 - Dec/03	398	**0,6398**	0,0000	0,0000	-	Yes
19 Singapore	Jan/95 - Dec/03	2.154	**-0,5717**	0,9999	0,9999	-	No
20 Sweden	Dec/95 - Dec/03	1.960	**-0,6319**	0,9999	0,9999	-	No
21 Switzerland	Jan/95 - Dec/03	2.183	**-0,5865**	0,9999	0,9999	-	No
22 **Turkey***	Ju/99 - Dec/03	878	**0,6324**	0,0000	0,0000	-	Yes
23 Thailand	Sep/95 - Dec/03	1.228	**-0,5810**	0,9999	0,9999	-	No
24 UK	Jan/97 - Dec/03	1.726	**-0,8535**	0,9999	0,9999	Yes - Negative	No
25 **Venezuela***	Jun/99 - Dec/02	629	**0,6884**	0,0005	0,0010	Yes - Positive	Yes

*Notes: (a) The p-values for each country qre calculated as follows: first, we estimate the AR coefficient of each country´s forward premium and country risk. Then, we simulate 3,000 pairs of series (one trying to mimic the properties of this country´s forward premium and the other series its country risk) with the estimated AR properties but we impose zero correlation on their stochastic shocks. (b) * denotes countries in which the country risk were measured by EMBI spread.*

When the series are cointegrated, we can ascertain that the estimated correlation coefficient is super-consistent, i.e., converges to the true value faster than it would if the series were stationary, so, in those cases the point estimate of the correlations on Table 1 are very reliable. We perform cointegration analysis on the series that were pairwise non-stationary[11]. In this case, the cousin risks phenomenon comes up when we do not reject the null hypothesis of cointegration between the two integrated series and the cointegration vector shows a positive relation between them. The detailed results of the Johanssen cointegration tests are presented on the Appendix[12] but are summarized on Table 1.

[11] Phillips-Perron unit root tests results are presented on the Appendix. Only 8 countries presented pairwise non-stationarity on country and currency risk premium: Argentina, Brazil, Czech Republic, Mexico, Poland, UK, South Africa and Venezuela.

[12] The coefficients reported on the appendix refers to the normalized cointegration vector $\begin{bmatrix} 1 & \beta \end{bmatrix} \begin{bmatrix} FP_t \\ CountryRisk \end{bmatrix} =$

We also constructed a distribution of the correlations by Monte Carlo Simulation and Bootstrap for each country under the null hypothesis that the forward premium and the country risk have no correlation and perform a non-parametric hypothesis test on each country's correlation coefficient. Our hypothesis test corrects for possibility of spurious correlation since in this exercise we take into account the high autoregressive properties (sometimes unit root) and the small sample properties of each country's series.

For each country, first, we regress the country risk on its own lag and the forward premium on its own lag as well. Next, we simulate 3,000 vectors $T_K x 2$, where T_K is the number of observations used to calculate the correlation of the FP and the country risk on country K, with the same AR coefficients of the country K's country risk and forward premium estimated in the previous stage. The stochastic term comes from a Monte Carlo simulation under the null hypothesis that the country risk and forward premium have zero correlation. The p-values of the estimated correlation on the distribution generated by this Monte Carlo Simulation for each country is presented in table 1. We implement the same exercise by Bootstrap and the results are also reported on table 1[13].

Based on these results, we propose a separation of the countries of our sample into two groups: (i) one composed of countries in which cousin risks phenomenon is observed and (ii) one composed of countries in which cousin risks phenomenon is NOT observed.

It is interesting to notice that both methodologies generate the same pattern of relationship. Among the countries that we could perform cointegration analysis, Argentina, Brazil, Mexico and Venezuela present positive cointegration relation and also have a positive correlation statistically different from zero. We could also perform cointegration analysis in Czech Republic, Poland, South Africa and UK and in all of them there is no presence of positive cointegration relationship and the correlations are not significantly positive.

Powell and Sturzenneger (2000) also studied Argentina, Brazil, Chile, Colombia, Mexico and Sweden, and their results are compatible with ours, except for the case of Chile. Indeed, Chile is a borderline country because although our hypothesis test can not reject that the correlation coefficient is significantly bigger than zero, we do not find any evidence of cointegration and its point estimate of the correlation is not too high (0.45). For these reasons, we follow Powell and Sturzenneger (2000) and classify Chile as not exhibiting cousin risks.

Table 2 summarizes our final proposed classification:

Table 2: Classification Proposed for the Countries Analysed

Cousin Risks Phenomenon	No Cousin Risks Phenomenon
Argentina, Brazil, Mexico, Peru, Phillipines, Russia, Turkey and Venezuela	Australia, Canada, Chile*, Colombia, South Korea, Indonesia, UK, Japan, Norway, New Zealand, Poland, Singapore, South Africa, Sweden, Switzerland and Thailand

** classification subject to robustness test*

0, so a negative β means a positive relation between the FP and the country risk.

[13] The matlab code for this simulation is available at www.econ.puc-rio.br/mgarcia/.

One of the main goals of our taxonomy is to permit the implementation of statistical tests to identify which variables are associated with the cousin risks phenomenon. Therefore, the classification is vital for the next section's results. For this reason, we implement robustness test of the next section's results by checking if the results would differ if Chile were classified as presenting the cousin risks phenomenon. The tests carried out in the appendix 4 do not point to significant changes in next section's results when the robustness checks are conducted.

4 Determinants of the Cousin Risks Phenomenon

4.1 Methodology and Data Description

Once identified which countries exhibit the cousin risks phenomenon, the next step is to apply a "DNA test" and determine what is linking them. In other words, what are the determinants of the risks co-movement? The most interesting feature of last section's results is that the cousin risks phenomenon does not constitute a rule among emerging countries. We will now investigate the cross-sectional dimension to try to uncover the cousin risks' determinants.

The discussion in Section 2 points to variables that could be responsible for the cousin risks so, in the present section, we will test if they are empirically associated with the presence of the phenomenon. This is done, first, by exploring these variable's statistical distribution among the different groups. Then, in last subsection, we present an econometric binary choice model[14].

4.2 Descriptive Statistics and Preliminary Tests

This subsection presents each country´s macroeconomic and financial data means from 1995 to 2002, almost the same time horizon we used in the last section to identify the phenomenon. The statistics are presented into three groups: (1) Countries that exhibit the cousin risks phenomenon; (2) countries that do not present the cousin risks phenomenon and (3) emerging market countries that do not exhibit the cousin risks phenomenon.

Our aim is to compare the distribution of each variable among the group of countries exhibiting cousin risks phenomenon and the group of countries not exhibiting cousin risks. In order to control for developed countries characteristics not captured in the sample (such as reputation), we also face the distribution of countries presenting the cousin risks phenomenon against the distribution of group of *emerging* countries not exhibiting cousin risks. In the following subsections we present tables with Kolmogorov-Smirnov

[14] The main data sources for this part of the paper are: The World Bank's World Development Indicators WDI, and IMF's International Financial Statistics IFS. Internal and external indebtedness data were obtained from each country's central bank, ministry of finance or statistics agency. Appendix 3 provides data sources and description of the variables.

tests, which determine if the distribution that originated two data sets differ significantly. To illustrate it graphically, we also plot the kernel densities[15] and the QQ plots.

4.2.1 Balance of Payment Variables

This subsection analyzes if a country's external 'health' (which is believed to be the main determinant of exchange rate expectations) is an important factor for the explanation of cousin risks phenomenon.

Table 3

Kolmogorov-Smirnov Test: Balance of Payment Variables	Cousin Risks Countries = Non-Cousin Risk (*p-value*)	Emerging Non-Cousin Risks (*p-value*)
Exports + Imports (% GDP)	0.2586	0.2929
Current Account Balance (% GDP)	0.9984	0.9438
Mean Import Tariff 1999 - 2000	0.0471	0.2523
International Reserves (% GDP)	0.6522	0.3874

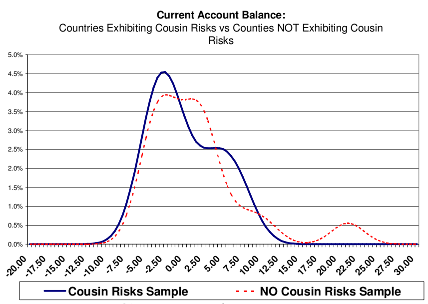

Fig. 5a: Current Account Densities

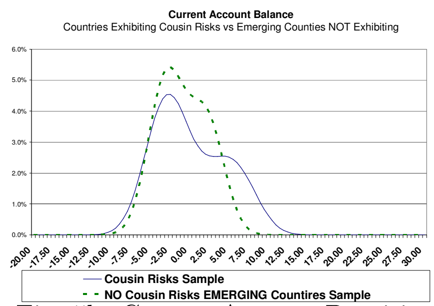

Fig. 5b: Current Account Densities

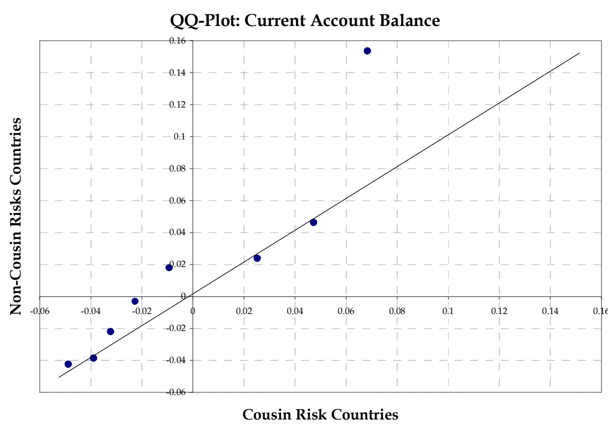

Fig. 6: QQ-Plot

The above data change only slightly from one group to another. This can be seen in figures 5a where we compare the densities of the sample of countries exhibiting the phenomenon against the sample of countries exhibiting the phenomenon and the samples is shown to be almost coincidental. This evidence does not change when we compare the sample of countries exhibiting cousin risks against the sample emerging countries not exhibiting it, as can be seen in figure 5b. The QQ-plots, where the quantile

[15] The bandwidth of this estimation is chosen as suggested by Silverman (1986).

of the cumulative densities of the two samples are compared, reinforce this observation. If they both came from the same distribution, the result would be points on the forty five degree slope line, what is precisely the case here. The same evidence stand out when we plotted the graphs on the other variables, so we omitted them. But the most important step is a formal statistical test and that what we turn to now.

Indeed, Kolmogorov-Smirnov test results, presented in Table 3, indicate that we cannot reject the null hypothesis that current account balance sample (%GDP) and exports plus imports sample (%GDP) among the group of countries exhibiting and not exhibiting cousin risks are statistically identical. The result is the same when we compare the countries exhibiting the phenomenon and *emerging* countries not exhibiting the phenomenon.

The only concern is import tariff. We reject the hypothesis that tariff import samples are identical among countries exhibiting and not exhibiting positive correlation between the country and the currency risk. However, when comparing only emerging markets we cannot reject the hypothesis that their sample are equal. This result is probably due to the fact that the sample of countries not exhibiting cousin risk is largely composed by developed countries that usually have lower import tariffs than emerging ones.

Thus the results of this section indicate that balance of payment indicators from countries that do exhibit the cousin risks do not differ significantly from countries in which the cousin risks phenomenon is not observed.

4.2.2 Solvency Variables

Since the country risk is a central variable to our study, government borrowing requirements and solvency variables are natural candidates to become the determinants of cousin risks. A possibility could be that countries with a fragile fiscal position exhibit a positive relation between country and currency risks.

Table 4

Kolmogorov-Smirnov Test: Solvency Variables	Cousin Risks Countries = Non-Cousin Risk (*p-value*)	Cousin Risks Countries = Emerging Non-Cousin Risks (*p-value*)
Total Public Debt (Internal + External %GDP)	0.1456	0.0214
Overall Budget Balance (% GDP)	0.4446	0.9438
Total External Debt (Government + Private % GDP)	0.5189	0.6852
Internal Government Debt (% GDP)	0.3541	0.6217

The Kolmogorov-Smirnov test results show that there is no distinction between these two groups in terms of the overall budget balance, total external debt (public + private) and government internal debt.

Nonetheless, governments of countries exhibiting cousin risks seem to be more indebted, as the Kolmogorov-Smirnov tests rejects the hypothesis that the sample from countries exhibiting cousin risks is equal to emerging countries not exhibiting cousin risks at 5% significant level. This result is weakened since we do not obtain a similar result when we compare cousin risks countries with the whole sample of countries not exhibiting cousin risks. Even so, we will further investigate it in next section, analyzing total indebtness jointly with other variables.

Therefore, solvency variables do not seem to determine the presence of the cousin risk phenomenon. We will advance on the analysis studying net effects in binary choice models in next section.

4.2.3 Financial Development and Currency Mismatch Variables

We now display the comparison of patterns of currency mismatch and financial development among the countries included in our sample.

Table 5

Kolmogorov-Smirnov Test: Currency Mismatch and Financial Deepening	Cousin Risks Countries = Non-Cousin Risk (p-value)	Cousin Risks Countries = Emerging Non-Cousin Risks (p-value)
Govt. External Debt - International Reserves (% GDP)	0.0002	0.0029
Gross Domestic Savings (% GDP)	0.1618	0.2929
Domestic credit to private sector (% GDP)	0.0004	0.0134
Market capitalization (% GDP)	0.0120	0.3380

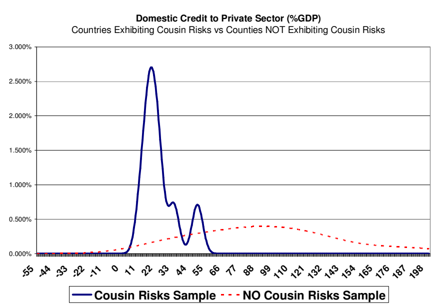

Fig. 7a: Domestic Credit Densities

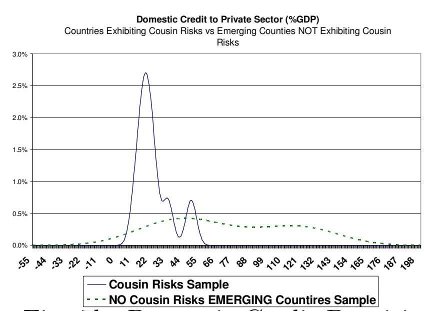

Fig. 7b: Domestic Credit Densities

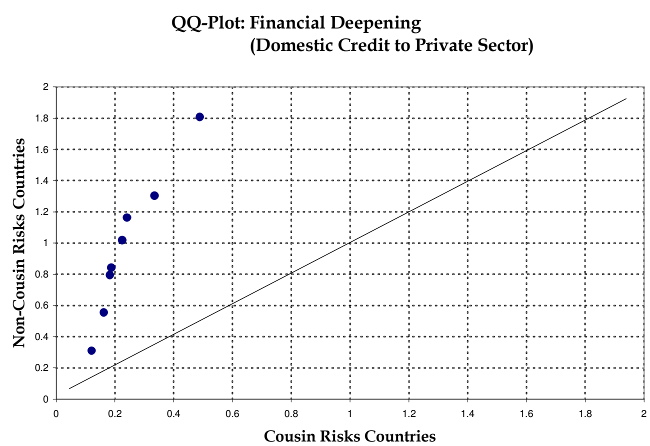

Fig. 8: QQ-Plot

Financial development is less intense in cousin risks countries. The Kolmogorov-Smirnov (KS) test rejects the hypothesis that these distributions are statistically equal: on the comparison of cousin risks countries with non-cousin risks countries it is rejected at the 1% significance level, on the comparison of cousin risks countries with non-cousin risks emerging countries at the 2% significance level. Indeed, these observations are reinforced by the location of the density distributions of cousin risk countries, to the left of the non-cousin risk countries distribution, as can be seen in figures 7a and 7b. The QQ-plot on figure 8 also illustrate that (as the points are far above the 45 degree line) domestic credit to private sector quantiles of the cousin risk sample (x-axis) are all lower than the non-cousin risk countries (y-axis).

The KS tests also highlight a striking difference between net exposure to exchange rate movements among the countries. The hypotheses that currency mismatch sample from cousin risks countries is equal to the ones from countries not presenting cousin risks (be they only the emerging ones or not) are rejected at the 1% significance level.

This indicates that the cousin risks phenomenon is associated with government's currency mismatch (external government debt minus international reserves) and the level of financial development (domestic credit for private sector). We now move to a framework that considers all factors at once: binary choice models.

4.3 Binary Choice Models

In this section we apply a binary choice model[16] using the same variables analyzed in last section. Following the taxonomy discussed in Section 3, the dependent variable assumes the value one for countries that exhibit the cousin risks phenomenon and zero for those that do not[17]. Results refer to Probit model output but the adoption of the Logit model did not qualitatively alter the results.

The explanatory variables are the same ones analyzed in the previous sections. Models contemplating different combinations of explanatory variables were estimated. Table 9 displays the models with only one explanatory variable, while Table 10 shows the results of multivariate analysis.

While working with Probit model, a positive (negative) coefficient significantly different from zero

[16] An alternative to binary choice models would be to use correlation as the dependent variable. Under such methodology, we apply the limited dependent variable models (such that the correlation is limited between -1 and +1) using cross-sectional data or we apply a more robust joint estimation of correlation, using the hierarchical linear model. However, in doing so, our already small sample would be tremendously reduced, thus harming the analysis. For example, in the case when the dependent variable is the correlation between the forward premium and the EMBI+ spread, only thirteen observation points can be included in the regression model. On the other hand, the adoption of the correlation between the forward premium and the CID would not reduce the sample size to the same extent, but the results would nonetheless be full of noises and less representative of investors' risk perception since CID measure is subject to regulatory and interventionist peculiarities of each country.

[17] A robustness test was carried out on our models, and the results are presented in Appendix, where we changed the classification of Chule. Major results don't change.

indicates that an increase the explanatory variable should increase (decreases) the probability of the country to exhibit the phenomenon.

Table 6: Probit Univariate Models

Dependent Variable: Cousin Risks (1 = exhibiting, 0= not exhibiting)
number of observations: 25

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8
constant	-1.318753	1.397137	2.101986	-1.311253	-0.776632	0.532655	-0.449209	-2.156276
p-value	*0.0025*	*0.2645*	*0.0024*	*0.0267*	*0.0283*	*0.466*	*0.1004*	*0.0037*
External Debt-Reserves (%GDP)	0.080538	-	-	-	-	-	-	-
p-value	*0.0006*	-	-	-	-	-	-	-
Savings (% GDP)	-	-0.080568	-	-	-	-	-	-
p-value	-	*0.1392*	-	-	-	-	-	-
Domestic Credit to Private Sector (% GDP)	-	-	-0.057412	-	-	-	-	-
p-value	-	-	*0.0026*	-	-	-	-	-
Total Debt (% PIB)	-	-	-	0.020622	-	-	-	-
p-value	-	-	-	*0.0791*	-	-	-	-
Overall Budget Balance (% PIB)	-	-	-	-	-0.164334	-	-	-
p-value	-	-	-	-	*0.1285*	-	-	-
Exports+Imports (% GDP)	-	-	-	-	-	-0.016657	-	-
p-value	-	-	-	-	-	*0.1611*	-	-
Current Account Balance (% GDP)	-	-	-	-	-	-	-0.032057	-
p-value	-	-	-	-	-	-	*0.552*	-
Mean Tariff Import	-	-	-	-	-	-	-	0.185956
p-value	-	-	-	-	-	-	-	*0.0083*
Schwartz criteria	0.86056	1.390352	0.78502	1.357284	1.387635	1.387417	1.494823	1.176535
McFadden's R2	0.518999	0.096429	0.57925	0.141852	0.098597	0.09877	0.013101	0.266973

Table 7: Probit Multivariate Models

Dependent Variable: Cousin Risks (1 = exhibiting, 0= not exhibiting)
number of observations: 25

	Model 9	Model 10	Model 11	Model 12	Model 13	Model 14
constant	-0.320948	1.909751	1.873102	0.898335	-1.628293	1.755246
p-value	*0.8459*	*0.0427*	*0.0301*	*0.5943*	*0.0046*	*0.0244*
External Debt-Reserves (%GDP)	0.083882	0.115055	0.106144	-	0.080556	-
p-value	*0.0016*	*0.0557*	*0.0022*	-	*0.001*	-
Savings (% GDP)	-0.047175	-	-	-	-	-
p-value	*0.5429*	-	-	-	-	-
Domestic Credit to Private Sector (% GDP)	-	-0.111958	-0.119056	-0.049203	-	-0.054909
p-value	-	*0.0426*	*0.0027*	*0.0222*	-	*0.0053*
Total Debt (% PIB)	-	-0.007048	-	-	-	-
p-value	-	*0.8495*	-	-	-	-
Overall Budget Balance (% PIB)	-		-	-	-0.134947	-0.104533
p-value	-		-	-	*0.3088*	*0.4006*
Exports+Imports (% GDP)	-		-	-	-	-
p-value	-		-	-	-	-
Current Account Balance (% GDP)	-		-	-	-	-
p-value	-		-	-	-	-
Mean Tariff Import	-		-	0.089394	-	-
p-value	-		-	*0.4508*	-	-
Schwartz criteria	0.979494	0.815347	0.660936	0.722177	0.961140	0.896633
McFadden's R2	0.526832	0.775597	0.780918	-	0.541471	0.592924

The results presented in Tables 6 and 7 support the findings in the last subsection. Univariate models, showed in Table 6 indicate that, at the 5% significance level, no solvency variable (Total debt or Fiscal result) significantly contributes to the explanation of the presence of the cousin risks phenomenon. Furthermore, the only external accounts variable that is significantly different from zero is the tariff level: the larger the mean import tariff, the larger the probability a country has of exhibiting cousin risks. Current account, as well as exports plus imports over GDP ratio, do not affect the country's probability of having the cousin risks phenomenon even at the 10% significance level. Gross domestic savings do not affect the probability of the cousin risks phenomenon occurrence even on 10% significance level. Currency mismatch and domestic credit for private sector are both statistically significant at 1% significant level. The higher the currency mismatch—defined as external debt minus international reserves—the higher the probability of the cousin risks phenomenon. Higher levels of financial development—calculated as credit for private sector—reduce the probability of a positive correlation between country and currency risk.

Multivariate models´ results are presented in Table 7. The most interesting feature is that government external debt minus international reserves and domestic credit to private sector are significantly different from zero in every model. Indeed, under both the Akaike and the Schwartz criteria, the best model is model 11 and according to McFadden´s R2 these two variables jointly explain more than 78% of the presence of cousin risks phenomenon. In all of the models, currency mismatch increases the probability and domestic credit to private sector reduces the probability of a country present cousin risks.

Models 13 and 14 show that the overall budget deficit, the gross domestic savings and the total government debt lost significance and do not help to explain the occurrence of cousin risk phenomenon when analyzed jointly with currency mismatch and financial deepening. Furthermore, while univariate models suggested that mean tariff import was important in determining the phenomenon, model 12 indicates that when we jointly analyze it with domestic credit, the tariff is no longer statistically significant.

The models are robust vis-à-vis the Chile's classification (see Appendix 9). Hence, we can conclude that the most important factors in determining the positive correlation between country risk and currency risk seems to be government currency mismatch and domestic credit to private sector.

Our interpretation of these statistical results is the following. Currency mismatch gives support to the causal link: given the currency mismatch, an increase in currency risk weakens balance sheets, thereby increasing country risk. Domestic credit to private sector gives support to the common generating factor link: when capital flows out, the impact on domestic physical investment and production will be stronger if only few domestic substitutes are available to provide financing. The larger the impact, the smaller the growth rate of the country and, hence, its ability to serve debt. Therefore, bad capital flows shocks should be associated with increases in both currency and country risk.

4.4 Adherence Analysis

We can see how well does the model fits the data for each country by checking its adherence. This is undertaken for model 11, the model with the best fit. Figure 11 displays the probability of the occurrence of the cousin risk phenomenon assigned by model 11. Ideally, countries that were classified as exhibiting the phenomenon (the red triangle ones) should be on the top of the graph, with 100% probability. The countries that were classified as not exhibiting the phenomenon should be on the bottom of the graph, with 0% probability. The evidence below suggests that we had a very nice fit and also allow us to identify the few countries in which the model fail to perform well.

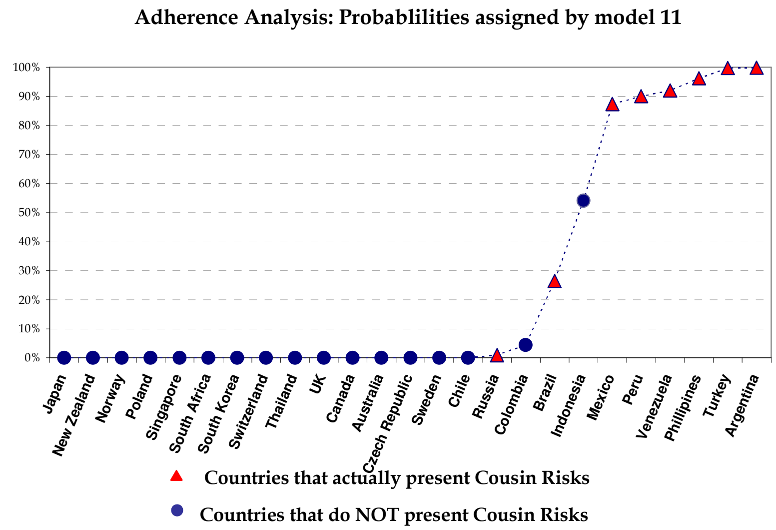

Fig. 9: Adherence Analysis

5 Conclusion

The positive correlation between country and currency risk premiums is referred to as cousin risks. Cousin risks is economically important because both risks are components of the domestic interest rate. Therefore, a country becomes more vulnerable to external shocks when these two risks are positively correlated, since negative shocks, as the reversal of capital flows, increase both risk premiums simultaneously while output is faltering. Cousin risks make interest rates higher and riskier, i.e., more volatile with covariances that amplify the deleterious effects of negative shocks to capital flows as sudden stops[18].

This paper focused on two main goals. The first one was to investigate how widespread the cousin risk phenomenon is, and the second goal was to identify the determinants of the correlation between these two risk premiums.

We identified that, among the countries in our sample (25 countries), Argentina, Brazil, Mexico, Russia, Peru, the Philippines, Turkey and Venezuela exhibit positive correlation between the country risk and the currency risk premiums. It is important to highlight that Chile, Colombia, South Korea, and South Africa do not exhibit positive correlation between these two risks premiums. Therefore, the cousin risks phenomenon is not omnipresent even among emerging markets.

In Section 4 we investigated the determinants of the cousin risks phenomenon. An interesting conclusion was that the sources of the cousin risks phenomenon are not the ones normally presented in the literature as determinants of country risk and currency risk premiums when they are independently analyzed. More specifically, the hypothesis that the balance of payments variables (which are believed

[18] Calvo, Izquierdo and Meija (2004)

to be the main sources of the currency risk premium) are responsible for the positive correlation between country risk and currency risk premiums is rejected. Based on our tests results, neither the level of indebtedness or surplus on fiscal accounts (which are the main determinants of the sovereign risk default) were accepted as being responsible for the cousin risks phenomenon.

Our empirical results indicate that the determinants of this phenomenon are:

1. Currency mismatch, measured as the difference between external government debt and international reserves (over GDP);

2. The level of financial deepening, measured by the credit to the private sector (over GDP);

Based on these results, we conjecture that when the government presents currency mismatch in its balance sheet, an increase in the expectation of exchange rate depreciation or an increase in exchange rate risk (both features are captured by forward premium) increase the perception of future government solvency condition, what, in turns, increases the sovereign credit risk. This would be the main channel through which currency risk would be associated with country risk.

The results are also an indication that cousin risks may be related to the original sin phenomenon (Eichengreen et al. (2002)). A country's inability to borrow in international financial markets in its own currency (original sin) causes a potential exchange rate mismatch. Eichengreen et al. (2002) holds that this can be harmful for those countries, and this paper claims that one of the main problems associated with the original sin is the occurrence of cousin risks. Indeed, cousin risks (which produce high and risky interest rates) and original sin appear to be different aspects of the same, more complex, phenomenon. If this is indeed the case, further examination of cousin risks may shed more light on the determinants of the original sin, as well as on the policy measures necessary to mitigate the deleterious effects of both phenomena.

Finally, high levels of credit to the private sector represent a substantial domestic supply of funds. The higher the level of financial deepening, the smaller the necessity of borrowing in international capital markets. In the event of reversal of capital flows, government and firms are able to resort to domestic finance if financial deepening is substantial. Therefore, in those events, investment would fall less, and so would GDP. On the other extreme, without the domestic credit market investment projects are interrupted and GDP suffers for long periods, thereby harming the countries ability to pay. This is what was called sudden stops.

In summary, we hypothesize that the two factors – exchange rate mismatch and financial deepening – generate the cousin risk phenomenon through two channels : (i) The exchange rate mismatch builds a causality link between currency risk and country risk through the balance sheet effects. In this channel, the reversal of capital flows would increase currency risk, which in turn, given balance sheet effects would increase country risk. (ii) The lack of deep domestic financial markets leverages the negative impacts of reversal of capital flows on investment and output, thereby harming the country's ability to pay. In this second channel, a third factor (reversal of capital flows) would simultaneously affect currency and country risk.

We see our characterization of the cousin risks phenomenon as a contribution to the growing literature on financial crisis affecting emerging markets. Theoretical and simulation models that generate interest rate data should take our results in consideration.

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6 Appendix

6.1 Unit root tests:

Phillips-Perron Unit Root Test

	FP	embi+ spread
	P-Value	P-Value
Argentina	0.995600317	1.000000000
Australia	0.980015014	0.000002161
Brazil	0.485506947	0.292443104
Canada	0.802150923	0.000033070
Chile	0.000000000	0.433621695
Colombia	0.000000000	0.231140390
Czech Republic	0.266613873	0.567461202
Indonesia	0.296228539	0.000000000
Japan	0.975541267	0.047035523
Mexico	0.424456258	0.356721968
New Zealand	0.815705860	0.000000000
Norway	0.771100293	0.000000000
Peru	0.000000743	0.678985596
Phillipines	0.000323509	0.220540304
Poland	0.540997625	0.109022705
Russia	0.000000000	0.000000059
Singapore	0.181343242	0.000000000
South Africa	0.480887331	0.745873025
South Korea	0.000020159	0.620878631
Sweden	0.000000000	0.000100000
Switzerland	0.897765617	0.001784714
Thailand	0.160629185	0.000000000
Turkey	0.000000000	0.514916414
UK	0.939619777	0.185104445
Venezuela	0.294404824	0.230181757

6.2 Cointegration tests:

Johanssen Cointegration Test Lags 0 0

	embi's spread vs. FP			
	Cointegration Test		Cointegration Vector	
	P-Value	Number of cointegration relation in HO	embi+	FP
Argentina	0.0000732	None	0.0033682	1.0000000
	0.0013602	At most 1	0.0007540	
Brazil	0.0008632	None	-0.0002153	1.0000000
	0.1114605	At most 1	0.0000351	
Czech Republic	0.3328062	None	0.0212048	1.0000000
	0.1232042	At most 1	0.0056713	
Venezuela	0.0001679	None	-0.0014367	1.0000000
	0.0274875	At most 1	0.0002322	
Mexico	0.0067277	None	-0.0002969	1.0000000
	0.1171373	At most 1	0.0000360	
Poland	0.0147073	None	-0.0006516	1.0000000
	0.0164341	At most 1	0.0002272	
UK	0.0000009	None	4.6790510	1.0000000
	0.9465100	At most 1	0.3716757	
South Africa	0.2717627	None	-0.0656215	1.0000000
	0.2248672	At most 1	0.0247351	

Johanssen Cointegration Test Lags 1 2

	embi's spread vs. FP			
	Cointegration Test		Cointegration Vector	
	P-Value	Number of cointegration relation in HO	embi+	FP
Argentina	0.0021377	None	-0.0011893	1.0000000
	0.0823005	At most 1	0.0002123	
Brazil	0.1901874	None	-0.0003252	1.0000000
	0.2374240	At most 1	0.0000683	
Czech Republic	0.4018867	None	0.0174022	1.0000000
	0.1025867	At most 1	0.0065526	
Venezuela	0.0057897	None	-0.0018123	1.0000000
	0.1108678	At most 1	0.0003025	
Mexico	0.0262723	None	-0.0002960	1.0000000
	0.0846490	At most 1	-0.0000400	
Poland	0.4996500	None	-0.0019121	1.0000000
	0.3704769	At most 1	0.0007107	
UK	0.0355407	None	4.5456027	1.0000000
	0.8787547	At most 1	0.5868052	
South Africa	0.2757992	None	-0.0638790	1.0000000
	0.2301085	At most 1	0.0243118	

6.3 Alternative correlation measures:

	FP-CID	FP-EMBI
	Daily	Daily
	Correlation	Correlation
1 S. Africa	**-0.7300**	**0.0612**
2 Argentina	**-0.4832**	**0.9267**
3 Brazil	**0.0590**	**0.7400**
4 Chile	**-0.8099**	**0.4572**
5 Colombia	**-0.4495**	**-0.4663**
6 S. Korea	**-0.7916**	**-0.6351**
7 Mexico	**0.4156**	**0.8609**
8 Peru	**0.3871**	**0.7162**
9 Phillipines	**0.2636**	**0.7521**
10 Poland	**0.1361**	**0.2535**
Correlation between the measures of correlation	0.60	

6.4 Data Sources:

The source of almost all financial markets historical quotation, such as spot and future exchange rate, interest rate swaps and treasury rates, is Boomberg. The excepltions are: (a) Brazil´s forward premium is calculated from dollar coupon "DDI" future rates and "DI" future rates and the source of these quotations is BMF (Brazilian Mercantile and Futures Exchange). (b) 1 year local interest rate in the following countries: Brazil (DI-pré), Mexico (TIIE 28), Colombia (CD 360) and Peru (deposit rate 1 year).

EMBI+ spread and EMBI GLOBAL spread are provided by JPMorgan.

Deffinition and Data Sources (except debt)

Variable	*Source:*
Exports of goods and services (% of GDP)	WDI - World Bank
Imports of goods and services (% of GDP)	WDI - World Bank
Current account balance (% of GDP)	WDI - World Bank
Simple Mean tarif	WDI - World Bank
Balance of Payments: Overall Balance	IFS - FMI
Gross international reserves (includes gold, current US$)	WDI - World Bank
External debt, total (DOD, current US$)	WDI - World Bank
Gross domestic savings (% of GDP)	WDI - World Bank
Government External Debt (% PIB)	Different each country (see next)
Government Internal Debt (% PIB)	Different each country (see next)
Overall budget balance, including grants (% of GDP)	WDI - Bco Mundial
Domestic credit to private sector (% of GDP)	WDI - Bco Mundial
Market capitalization of listed companies (% of GDP)	WDI - Bco Mundial

Debt Data Sources

	Internal Debt	External Debt
Argentina	Ministerio de Economia y Producción	Ministerio de Economia y Producción
Australia	OECD	RBA - Reserve Bank of Australia
Brazil	BCB - Banco Central do Brasil	BCB - Banco Central do Brasil
Canada	OECD	SDSS IMF
Chile	Ministério da fazenda do chile (Deuda del Gobierno Central)	Ministério da fazenda do chile (Deuda del Gobierno Central)
Colombia	Banco de la República - Colômbia	Banco de la República - Colômbia
Czech Rep.	IFS - IMF	IFS - IMF
Indonesia	World Bank	World Bank
Japan	OECD	Ministry of Finance - Japan
Mexico	Secretaria de Hacienda - Mexico	Secretaria de Hacienda - Mexico
New Zealand	OECD	Reserve Bank of New Zealand
Norway	SDSS IMF	SDSS IMF
Peru	Banco Central de Reserva del Perú	Ministerio de Economia y Finanzas
Phillipines	Department of Economic Research - Bangko Sentral ng Pilipinas	Department of Economic Research - Bangko Sentral ng Pilipinas
Poland	IFS - IMF	IFS - IMF
Russia	Ministry of Finance	Ministry of Finance
Singapore	Singapore Department of Statistics	Singapore Department of Statistics
South Africa	IFS - IMF	IFS - IMF
South Korea	Ministry of Finance - Korea	Bank of Korea
Sweden	Statistiska centralbyrán	Statistiska centralbyrán
Switzerland	-	-
Thailand	IFS - IMF	IFS - IMF
Turkey	Central Bank of the Republic of Turkey	Central Bank of the Republic of Turkey
UK	OECD	-
Venezuela	The Ministry of Finance	The Ministry of Finance

6.5 Models and tests with Chile classified as cousin risk country:

Kolmogorov-Smirnov Test: Balance of Payment Variables	**Cousin Risks Countries =** **Non-Cousin Risk** *(p-value)*	**Emerging Non-Cousin Risks** *(p-value)*
Exports + Imports (% GDP)	0.1367	0.1400
Current Account Balance (% GDP)	0.8315	0.9836
Mean Import Tariff 1999 - 2000	0.0329	0.5327
International Reserves (% GDP)	0.5412	0.7249
Total Public Debt (Internal + External %GDP)	0.5506	0.1400
Overall Budget Balance (% GDP)	0.7490	0.8938
Total External Debt (Government + Private % GDP)	0.2763	0.5327
External Government Debt (% GDP)	0.0100	0.0224
Internal Government Debt (% GDP)	0.2540	0.8938
Govt. External Debt - International Reserves (% GDP)	0.0018	0.0224
Gross Domestic Savings (% GDP)	0.3345	0.1400
Domestic credit to private sector (% GDP)	0.0012	0.0435
Market capitalization (% GDP)	0.0511	0.8938

Probit Univariate Models

Dependent Variable: Cousin Risks (1=Presenting, 0=Not presenting)
number of observations: 25

	Model 1	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**	**Model 7**	**Model 8**
constant	-0.685837776	1.360195864	1.816484019	-0.924017084	-0.59628719	0.62644411	-0.33478807	-2.1264817
p-value	*0.0606*	*0.206853213*	*0.005748296*	*0.02405861*	*0.050146512*	*0.382862093*	*0.198697462*	*0.001042083*
External Debt-Reserves (%GDP)	0.046850804	-	-	-	-	-	-	-
p-value	*0.0169*	-	-	-	-	-	-	-
Savings (% GDP)	-	-0.073595921	-	-	-	-	-	-
p-value	-	*0.100873966*	-	-	-	-	-	-
Domestic Credit to Private Sector (% GDP)	-	-	-0.042533996	-	-	-	-	-
p-value	-	-	*0.000495571*	-	-	-	-	-
Total Debt (% PIB)	-	-	-	0.014771431	-	-	-	-
p-value	-	-	-	*0.050027866*	-	-	-	-
Overall Budget Balance (% PIB)	-	-	-	-	-0.13458361	-	-	-
p-value	-	-	-	-	*0.078165038*	-	-	-
Exports+Imports (% GDP)	-	-	-	-	-	-0.01620215	-	-
p-value	-	-	-	-	-	*0.170592538*	-	-
Current Account Balance (% GDP)	-	-	-	-	-	-	-0.04667611	-
p-value	-	-	-	-	-	-	*0.305229593*	-
Mean Tariff Import	-	-	-	-	-	-	-	0.19769194
p-value	-	-	-	-	-	-	-	*0.00640766*
Schwartz criteria	1.161322345	1.452017281	0.890716702	1.474008485	1.471393074	1.38741708	1.53006311	1.18056845
McFadden's R2	0.308396762	0.08595504	0.515466021	0.086126166	0.071128553	0.09877009	0.02623385	0.29366951

Probit Multivariate Models

Dependent Variable: Cousin Risks (1=Presenting, 0=Not presenting)
number of observations: 25

	Model 9	**Model 10**	**Model 11**	**Model 12**	**Model 13**	**Model 14**
constant	-0.23292992	1.314544955	1.311995	-0.205439069	-0.783989595	1.653985725
p-value	*0.848909608*	*0.094869485*	*0.104682113*	*0.85093611*	*0.069647695*	*0.038865045*
External Debt-Reserves (%GDP)	0.045522465	0.020077341	0.019727	-	0.044759871	-
p-value	*0.031523995*	*0.386581423*	*0.327883684*	-	*0.021913755*	-
Savings (% GDP)	-0.019461132	-	-	-	-	-
p-value	*0.695825764*	-	-	-	-	-
Domestic Credit to Private Sector (% GDP)	-	-0.035709787	-0.035819617	-0.031595297	-	-0.040769326
p-value	-	*0.003234299*	*0.00230323*	*0.001081259*	-	*0.001657743*
Total Debt (% PIB)	-	-0.000569151	-	-	-	-
p-value	-	*0.960926186*	-	-	-	-
Overall Budget Balance (% PIB)	-	-	-	-	-0.057332939	-0.042542086
p-value	-	-	-	-	*0.528991694*	*0.708578544*
Exports+Imports (% GDP)	-	-	-	-	-	-
p-value	-	-	-	-	-	-
Current Account Balance (% GDP)	-	-	-	-	-	-
p-value	-	-	-	-	-	-
Mean Tariff Import	-	-	-	0.16098981	-	-
p-value	-	-	-	*0.044403712*	-	-
Schwartz criteria	1.286666	1.144395	0.976422	0.962757421	1.281479	1.016191
McFadden's R2	0.311007483	0.535403976	0.548408217	0.558864196	0.314976283	0.517976743

Nesta tese foram desenvolvidos três ensaios nos quais foram utilizados arcabouços de finanças com o objetivo de estudar três questões de macroeconomia aplicada. E nos apêndices, que terminam aqui, apresentamos os artigos, que foram escritos originalmente em inglês, nos quais foram apresentadas as derivações matemáticas dos modelos e efetuados diversos testes de robustez dos resultados empíricos apresentados ao longo do corpo da tese.