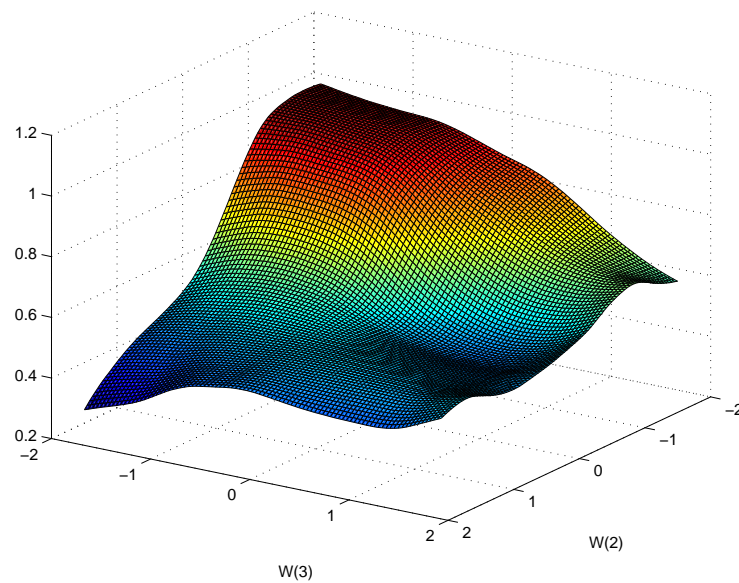


# Revising Arguments for the Modern Policy Framework - an Empirical Study

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Für meine Familie

# Vorwort

Die vorliegende Arbeit entstand während meiner Tätigkeit als wissenschaftlicher Mitarbeiter am Institut für Statistik und Ökonometrie der Christian-Albrechts-Universität zu Kiel. Mein besonderer Dank gilt meinem Doktorvater Professor Dr. Helmut Herwartz, dessen erstklassige Betreuung entscheidend zum Gelingen dieser Arbeit beigetragen hat. Herrn Professor Dr. Stefan Reitz danke ich herzlich für die Übernahme des Zweitgutachtens.

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Kiel, am 14. November 2011

Jan Roestel

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# Chapter 1

## Introduction

Pursuing a low inflation policy in an environment of open capital accounts and flexible exchange rates is a basic strategy that is widely adopted by industrialized and emerging economies. The general appeal of this framework can be summarized by few basic arguments. Namely, the fact that so many economies aim to keep their inflation rates low and stable reflects the view that economic growth and stability can be achieved more easily in this environment. Openness of capital accounts is regarded as indispensable since it promises significant welfare gains from improved capital allocation. Lastly, policymakers prefer free floating exchange rates to alternative policies of imperfect flexibility to be better able to pursue a domestic monetary policy under globalized capital markets.<sup>1</sup> While the empirical investigation of such claims has important implications for the most fundamental decisions of monetary policy, extant studies suffer from several conceptual shortcomings that might complicate straightforward conclusions. The goal of this work is to reinvestigate the aforementioned claims from an empirical point of view, taking into account associated pitfalls in related research. In the following four main chapters, each issue will be addressed separately. The interrelationship between inflation, inflation uncertainty, economic growth and stability will be studied in Chapter 2, the current state of capital market integration in Chapter 3 and monetary independence in Chapter 4 and 5. Chapter 3 corresponds to the article of Herwartz and Roestel (2011) as published by John Wiley and Sons. Chapter 5 is the article of Herwartz and Roestel (2009) as published by De Gruyter. Subsequently, I will discuss each issue in more detail and illustrate in which respect existing empirical studies might give rise to misleading conclusions. By reference to this literature, I will highlight the contribution of this work and summarize the results.

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<sup>1</sup>Apart from the benefit of monetary independence, a flexible foreign exchange rate further offers the opportunity to reestablish cross country competitiveness in presence of sticky prices and wages at low cost. In this study, I will concentrate on the aspect of monetary independence, while the aforementioned issue of exchange rate pass through will be addressed in future research.

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The view that inflation and inflation uncertainty tend to impair economic growth and stability will be empirically confirmed in Chapter 2. For this purpose, the joint interrelationship between inflation, output growth and respective uncertainties will be estimated and discussed for 34 developed and emerging economies. Policymakers are reluctant to accept higher inflation since this potentially triggers inflation uncertainty, which might incur costs in terms of output uncertainty and lower economic growth. However, the financial crisis has initiated a general reconsideration of the paradigm of low inflation policies. Nowadays, increasingly many economists suggest to raise inflation levels moderately, stressing that this might increase the power of the interest rate instrument. For instance, Blanchard et al. (2010) advocates that inflation targeting strategies should be modified in the sense that inflation expectations are well anchored, but on a moderate level, 4%, say. In this context, the potential economic costs of such a policy are often condensed to be negligible. However, several assumptions underlie such arguments. Namely, that i) one may separate inflation levels and inflation uncertainty, ii) inflation uncertainty rather than inflation is detrimental to growth iii) inflation and/or inflation uncertainty does not trigger output volatility, or, alternatively, iv) output volatility is at least not harmful to output growth. For such questions, it is of importance to understand the joint interaction of inflation, its uncertainty, economic growth and stability that tends to prevail in modern-world economies with low to moderate inflation rates. However, studies that assess these joint interrelationships such as Grier and Perry (2000), Fountas and Karanasos (2007) or Bredin and Fountas (2009) often provide conflicting evidence over distinct economies, while relying on quite historic data sets. These tend to comprise information obtained during the eras of Bretton Woods, oil crises and the Great Moderation, raising the question if such information is useful for current policy issues. In contrast, we will assess the simultaneous interaction between inflation, inflation uncertainty, output growth and volatility for 34 economies during the current era of low inflation policies. Obviously, this requires the estimation of large scale econometric models under use of very short sampling periods. However, we will argue that selective inference is possible in the framework of mean group estimation. After introducing the econometric methodology we will discuss the empirical findings derived from cross sectional features of country specific transmission parameter estimates. Thereby, we will put a particular effort on highlighting the role of monetary performance, providing characteristic transmission patterns for economies with low, moderate and higher inflation rates. One main insight of the analysis is that both inflation and inflation uncertainty tend to reduce output growth for the vast majority of investigated economies and, moreover, this

link cannot be rejected to be equally strong in economies with low, intermediate and higher inflation rates. Hence, it is unlikely that one could target inflation rates at intermediate levels (rather than at low levels) at no cost, and hence, keeping inflation rates low and stable is an attractive policy option.

Turning to the the argument that open capital accounts give rise to investment allocation efficiency gains, Chapter 3 provides empirical evidence that capital market integration has likely prevailed for a subset of G7 countries during the most recent decade. The vast majority of economies are reluctant to impose barriers to international capital flows despite the fact that they have been severely affected by international financial turmoils during the past two decades. This shows that economists still agree on the overall benefit of capital market openness. Recent studies such as Laeven (2003), Galindo, Schiantarelli and Weiss (2007) or Abiad, Oomes and Ueda (2007) suggest that capital market openness is accompanied by higher investment allocation efficiency, smoothes consumption growth (Bekaert, Harvey and Lundblad 2006) and fosters economic activity (Rousseau 2003, Bekaert, Harvey and Lundblad 2005, Bekaert, Harvey, Lundblad and Siegel 2007).<sup>2</sup> To some extent, however, all of these benefits reflect a consequence of higher investment allocation efficiency. The most direct approaches to test for capital allocation efficiency are based on the law of one price, which implies that capital can be raised at equal global cost. An intuitive way to diagnose this form of efficiency is to test for the convergence of international real interest rates.<sup>3</sup> In presence of integrated world markets, both the return on physical and financial assets should tend to converge across economies. While this proposition has received considerable attention in the empirical literature, evidence for full convergence among international real interest rates is rather mixed. I will argue that extant studies are subject to particular deficiencies of real interest rate measurement which do not apply in case that capital market real interest rates are measured by means of long term inflation indexed interest rates. Considering the most recent decade, we will test if capital can be raised at equal cost for a subset of the G7 economies, focussing on the real cost of long term external capital obtained from inflation indexed government bonds. After providing a brief discussion on the economic theory and the considered hypotheses we will discuss the methodologies that we use to infer on real interest rate convergence properties. Then, we will present the empirical results

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<sup>2</sup>In contrast, Prasad, Rogoff, Wei and Kose (2003) document a negative impact on growth.

<sup>3</sup>This concept does not rely on potentially misspecified asset pricing models commonly used in international finance. Here, asset pricing models are used to uncover if the cost of equity capital is determined with respect to global or local risks (see e.g. Bekaert and Harvey 1995, Hardouvelis et al. 2006, Carrieri, Errunza and Hogan 2009).

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that suggest complete long run convergence of real interest rates. In addition, we will provide empirical evidence suggesting that particular deficiencies of real interest rate measurement could give rise to a spurious rejection of capital market integration. We conclude that international capital markets are integrated nowadays, and hence, maintaining open capital accounts is a beneficial policy option.

Supporting the view that flexible exchange rates give rise to a higher degree of monetary independence compared to less flexible arrangements, Chapter 4 provides an analysis of domestic interest rate dependence under distinct degrees of exchange rate flexibility. According to Mundell (1963), monetary authorities should be able to pursue an independent monetary policy in presence of open capital accounts if spot exchange rates are allowed to move freely. Hence, small economies are only able to follow monetary strategies such as inflation targeting if they accept some exposure to exchange rate volatility and, hence, a potential impediment for cross border trade and investment. The main points in Mundell (1963) are summarized in the 'macroeconomic trilemma' (Obstfeld, Shambaugh and Taylor 2005). Namely, that only two among the following three fundamental aims of monetary policy can be achieved jointly: Capital account openness, fixed exchange rates and monetary independence. Firstly, an economy with an open capital account and pegged exchange rates cannot conduct an independent monetary policy. Its central bank would have to react to the prevalence of excess supply/demand for local currency whenever international investors are not indifferent on the cross country interest rate spreads that emerge in the conduct of domestic monetary policy. Such interventions on foreign exchange rate markets, however, give rise to balance-of-payments-induced changes in the money supply that tend to offset the initial interest rate manipulation. Therefore, under pegged exchange rates and open capital markets, the money supply and the interest rate are no longer available means of a proactive domestic monetary policy. Secondly, small open economies can only set domestic interest rates independently from foreign ones if the price of the domestic currency is flexible enough to reestablish the indifference of investors on the resulting cross country nominal interest rate spreads. Lastly, an economy that intends to pursue a domestic monetary policy under fixed exchange rates cannot allow for open capital accounts. Interest rates and exchange rates can only be chosen independently from each other if international capital is immobile. Noting that, in general, policymakers are reluctant to accept high exchange rate volatility, one might be interested in the question if flexibility of exchange rates must be perfect to guarantee a sufficient degree of monetary independence. For instance, if monetary authorities only allow for graduate adjustments, linkages among

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interest rates might be sufficiently weak to establish the possibility to address domestic goals at least temporarily. Hence, apart from extreme cases such as free floating or fixed exchange rates, the macroeconomic trilemma in fact suggests a continuous tradeoff among interest rate independence and exchange rate stability. In this respect, existing studies only provide limited insights as they tend to refer to highly stylized classifications of FX rate flexibility. To evaluate the argument if floating exchange rates allow for a higher independence of monetary policy than less flexible arrangements, I assess the macroeconomic trilemma as generally and flexibly as possible by considering interest rate dependence as a nonlinear function of exchange rate variability and capital mobility. This issue is analyzed in a semiparametric panel framework with functional coefficients (Cai, Fan and Yao 2000, Herwartz and Xu 2009). Empirical evidence suggests that nominal interest rate linkages tend to increase gradually with exchange rate stability and capital market openness, while perfect FX flexibility does not appear indispensable for some degree of monetary independence. Formally testing the implications of the macroeconomic trilemma, I strongly reject invariance of interest rate transmission with respect to capital mobility and FX rate flexibility. After a further motivation, we will introduce the empirical model, which is estimated by means of both conventional and semiparametric panel techniques. Furthermore, the employed state variables for measuring exchange rate flexibility and capital mobility are introduced and motivated. Subsequently, we will present and discuss the empirical results suggesting that flexible exchange rates give rise to a higher degree of monetary independence as compared to less flexible arrangements.

In Chapter 5, we will provide an alternative empirical assessment of monetary dependence under flexible exchange rates for a subset of G7 economies, namely Canada, France (assumed to represent the EMU), the UK and the US. Rather than estimating the comovements of short term variables such as short term nominal interest rates or inflation rates, I consider the empirical dependence of ex-ante long run inflation rates. These are measured in terms of international breakeven inflation rates. Proceeding this way one may avoid common pitfalls related to the measurement of independence by means of comovements among short term interest or inflation rates as in Eun and Jeong (1999), Yang et al. (2006), Frankel et al. (2004) or Obstfeld et al. (2005). These are related to the fact that nominal comovements might reflect central banks to independently choose a similar stance of monetary policy when facing common shocks or synchronized business cycles. Given that ten year ahead inflation compensation should predominantly reflect expectations regarding long run monetary policy, it is unlikely that such spurious patterns of dependence dominate nominal linkages observed across

## Chapter 1 Introduction

countries. I analyze nominal linkages both in terms of testing Granger causality in the framework of a VAR model and in terms of contemporaneous correlation which is interpreted to reflect both common shocks and nominal linkage. I find that in particular long run inflation prospects in Canada vary with expectations concerning long run US inflation, thereby failing to fully reflect the domestic inflation target. Similarly, also inflation expectations in the Euro area and the UK are subjected to common shocks, but empirical correlations vary on a lower scale. Thus, in comparison with the Bank of Canada, the Bank of England is likely better able to target domestic long run inflation. More specifically, empirical findings suggest that Canadian and US rate correlations are relatively high if *i*) the US conducts an expansionary monetary policy, *ii*) US inflation uncertainty is high or *iii*) if the Canadian-US exchange rate rate volatility is low. Similarly, UK and EMU linkage appears to increase if (long term) inflation expectations in the Euro area become relatively uncertain. Nominal linkage between the UK and the US rises in response to US inflation uncertainty and volatile UK FX reserves that likely indicate FX market interventions of the Bank of England. It therefore appears that future inflation prospects in small floating economies are influenced by the monetary policy of its respective large trading partner. After a further motivation, we will introduce the data and sketch the liquidity adjustment of breakeven inflation. Subsequently, a VAR analysis for international breakeven rates is performed. As a viable long run relationship cannot be detected, monthly realized correlations of breakeven rate changes are determined and conditioned on measurable economic variables. Then, the findings and its implications for monetary policy are discussed.

## Chapter 2

# Inflation, inflation uncertainty and its impact on real activity

Understanding the interaction of inflation, inflation uncertainty, economic growth and stability is an important aspect for the setup of theoretical models and the conduct of monetary policy. However, empirical and theoretical contributions especially disagree with regard to the influence of uncertainty on the joint determination of the conditional means of inflation and output. Empirical studies find conflicting evidence in particular if distinct economies are analyzed (Grier and Perry 2000, Fountas and Karanasos, 2006, Bredin and Fountas 2009, Caporale and Kontonikas 2009). Arguments from theoretical contributions suggest several influences on particular causal relations which might be economy-specific. For example, Cukierman and Meltzer (1986), Devereux (1989) or Ball (1992) formalize the interaction between inflation uncertainty and inflation as a matter of (expected) interventions of monetary policy. Examples of empirical studies that address such interrelations are Kontonikas (2004) or Caporale and Kontonikas (2009). These studies document that the adoption of credible and transparent strategies such as inflation targeting (IT) can influence the interaction among inflation and inflation uncertainty (IU). Similarly, Fountas et al. (2004) or Gillman et al. (2004) discuss dependence of the relation between IU and output on the level of inflation. The inflation rate is also supposed to influence relations between IU and the uncertainty about output (OU) in contributions by Taylor (1994) and Fuhrer (1997), as well as the trade-off among economic growth and inflation (Ball, Mankiw, and Romer 1988).

In case that such relations in fact depend on country specific characteristics related to monetary policy, the ambivalence of empirical findings is well in line with the predictions of aforementioned theoretical approaches. Conflicting empirical evidence could reflect that monetary policy objectives tend to change over economies and time.

During the recent decades, however, the hegemony of inflation targeting strategies has possibly led to some degree of harmonization regarding the policy schemes adopted in many economies. If particular effects are relevant within the current 'common sense' monetary policy framework, one might expect to find them in similar form across distinct types of economies though other economic characteristics might be markedly distinct.

The aim of this study is to empirically identify common features in such interrelationships across a large range of distinct economies and a recent sample period. Arguing that empirical investigations on causality might be conducted as tests on whether particular linkages among variables exist reliably across distinct types of economies, we estimate uniform bivariate GARCH-in-mean VAR models (GARCH-M for short)<sup>1</sup> for 34 industrialized and emerging economies. For inference, we use the cross sectional information obtained from country specific parameter estimates.<sup>2</sup> Extant studies discuss causality by examining if particular patterns of temporal ordering tend to occur systematically in individual economies. We instead regard a relation as economically relevant if respective point estimates of intertemporal effects can be found in similar (nonzero) magnitudes for distinct economies. We investigate the validity of particular causalities as hypothesized by Okun (1971), Taylor (1994), Friedman (1977), Cukierman and Meltzer (1986), Devereux (1989), Black (1987) or Pindyck (1991), but also document other systematic linkages that are less extensively discussed in the theoretical literature. Furthermore, we investigate to which extent monetary stability influences particular causalities by formally comparing characteristic transmission dynamics across economies with low, medium or higher inflation rates.

By relying on the cross sectional variation of country specific parameter estimates for inferential purposes, we document several regularities asserted by economic theory in an international context. Our approach allows to conduct inference regarding the joint interactions among uncertainty variables and conditional means, along with explicit volatility spillover effects. Our analysis is based on a rather large cross section of 34 economies and is, furthermore, focused explicitly on the most recent policy era. In related studies, evidence is mostly country specific, while the large amount of information required for inference in such general models stems from a plentitude of

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<sup>1</sup>Such specifications are a widely used means of quantifying and testing respective causal effects (Grier and Perry 2000, Bredin and Fountas 2009, inter alia).

<sup>2</sup>Thereby we acknowledge arguments put forward e.g. in Grier and Grier (2006) who note that evidence from developing countries might be a useful means to put findings from industrialized economies into context.



time observations, typically 40-50 years of monthly data. Such settings restrain the analysis to smaller cross section dimensions (such as the G7) where data regarding longer sampling periods is available. Moreover, during such long periods, the dynamics of inflation and output in many economies might have been subject to structural change (Evans and Wachtel 1993, Garcia and Perron 1996, Caporale and Kontonikas 2009). Proceeding from the assumption of structural stability, however, might lead to unclear implications for the informative value of estimates and hypothesis testing.<sup>3</sup>

For the considered economies, we document a markedly negative effect of inflation on output growth. Evidence for the opposite causal direction running from output to higher inflation is also found but the effect is of considerably smaller magnitude. Moreover, also inflation uncertainty adversely affects economic activity. Apart from this linkage, GARCH-implied quantifications of uncertainty do not systematically influence output growth or inflation levels. Spillover effects among volatilities of output and inflation are potentially present, but are, however, of rather small magnitude. Formally comparing if particular effects systematically differ across economies, we find that real effects of inflation appear especially strong in countries with low inflation rates. Surprisingly, the impact of uncertainties appears to be unaffected by monetary performance. The remainder of the paper is organized as follows. In Section 2, we introduce our econometric model. In Section 3, we discuss estimation results and provide model diagnostics along with robustness checks. Section 4 concludes.

## **2.1 Empirical Framework**

In this Section, we describe the econometric model and the inferential strategy of testing the economic relevance of causality hypotheses as suggested by economic theory. Firstly, we introduce the data. The GARCH-M model, which jointly expresses first- and second order dynamics in inflation and output is described subsequently. After introducing the model, the methodology used to test for causal effects by consideration of cross-sectional variation is described.

An important aspect of this study is that our data set fundamentally differs from those that are typically considered for joint assessments of causal effects among inflation, output and respective uncertainties. The consideration of a data set which is focused on the most current time periods and a large cross section might increase the relevance of results for monetary policy. Moreover, it might also be beneficial in

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<sup>3</sup>Accounting for structural shifts in heavily parameterized models is problematic due to the inefficiency of resulting estimates.

terms of the reliability of empirical findings. Data sets which include long series of historical observations are likely to incorporate distinct macroeconomic regimes (Bruno and Easterly 1995, Giordani and Söderlind 2003). If changes in regimes are not taken into account appropriately, this may lead to biased conclusions on causal linkages. The issue of structural instability is usually addressed by means of introducing more flexible specifications which allow for changes in the parametrization of the VAR or GARCH process (e.g. Evans and Wachtel 1993, Kontonikas 2004, Bredin and Fountas 2009). Proceeding in this way, however, requires to determine the precise timing of structural changes. This problem has to be either resolved by means of testing procedures (Bai and Perron 2003) or by introduction of a priori knowledge about the timing of influential events into the modeling process. Both methods, however, are of limited applicability if a large cross sectional data set is examined like in the case of this study. Moreover, in case that many model parameters change, one might ask if this approach promises a significant improvement compared to reducing the size of the sample.

For these reasons we seek to focus on the most recent time between 1990 and 2010. Regarding the widespread adoption of inflation targeting strategies and the recognition of excess (and instable) inflation as detrimental to economic growth during this time, the last two decades might be labeled as the era of inflation targeting. In focusing on this inflation targeting period, we also circumvent the problem of determining the timing of potential changes in the parametrization of the interrelation between first and second order moment dynamics.

### 2.1.1 The data

The observed variables that we consider are annualized rates of inflation  $\pi_t = 1200 \times \log(CPI_t/CPI_{t-1})$  and output growth  $y_t = 1200 \times \log(IP_t/IP_{t-1})$  where  $CPI_t$  and  $IP_t$  denote the consumer price and industrial production index, respectively. All series are seasonally adjusted by means of the Tramo/Seats method. We incorporate output growth not only for reasons of comparability to extant studies like Bernanke and Blinder (1992), Elder (2004) or Fountas and Karanasos (2004). Growth rates have the advantage of avoiding to rely on estimated quantities as potential output, which is necessary for measures of the output gap. The oil price growth rates  $oil_{it}$  are obtained for crude oil of the sort West Texas Intermediate priced in terms of domestic currencies, that is  $oil_{it} = 1200 \times \log(WTI_{it}/WTI_{i,t-1})$ . The data set covers monthly observations for the time  $t = 1990M1$  until  $t = 2010M1$  and 34 economies. The cross section of economies is enlisted in Table 2.1.

To examine the time series characteristics of inflation and output growth rates we firstly resort to the application of heteroskedasticity consistent ADF unit root tests. In case that country specific steady state inflation does not substantially change over the considered sampling period one would expect inflation rates to be stationary. Furthermore, we test for the presence of multivariate ARCH effects in order to check if the modeling of conditional second moments is justified for these particular data series. Table 2.1 reports the outcomes of ADF unit root tests for the inflation series in each country while ADF statistics for output growth rates, throughout signalling stationarity, are not listed explicitly. Inspection of ADF diagnostics suggests that only 6 out of 34 countries feature nonstationary inflation rates for the considered sampling period, namely Argentina, Brazil, Columbia, Hungary, Peru and Poland. In these economies, inflation levels have been rather high during recent decades and their reduction might still be at an early stage compared with the rest of the economies under consideration. To summarize, it appears unnecessary to transform the inflation and output growth series in a particular way prior to setting up the econometric model. The overall evidence on stationary inflation rates likely reflects that most of the countries have successfully pursued a credible low inflation policy during the past two decades.

Comparison of these results to unit root tests for inflation rates obtained for a longer period of 1980-2010, say, reveals that nonstationarity can be rejected in a substantially smaller number of cases. It seems that a focus on the two recent decades is beneficial as this period is characterized by a larger degree of structural stability. Table 2.1 further shows multivariate ARCH-LM test diagnostics obtained for the residuals of the model (2.2). Respective diagnostics indicate significant ARCH effects in the nonstandardized model innovations for the vast majority of economies. In any case, a potential absence of ARCH effects can be accommodated by our model as this corresponds to a special case of equation (2.1). Therefore, the inclusion of potentially homoscedastic inflation and output growth series may provide useful information regarding the overall (in)significance of causality parameters.

### **2.1.2 The bivariate GARCH-in-mean model**

Inflation and output uncertainty may be defined and estimated in several ways (Lahiri and Liu 2005). Specifications based on the (G)ARCH methodology (Bollerslev 1986, Baillie 1996) have become one of the most widespread means of measuring such unobserved variables. Engle (1982) and Ball and Cecchetti (1990) emphasize that it is in particular the unpredicted component of overall variation which might be

Table 2.1: Test diagnostics: Unit roots, Multivariate ARCH

	ADF $\pi$	LM		ADF $\pi$	LM		ADF $\pi$	LM
Argentina (AG)	-7.636	144.683	Hungary (HN)	-1.85	103.713	Norway (NW)	-11.73	60.530
	(0.49)	(0.00)		(0.61)	(0.00)		(0.00)	(0.00)
Austria (AT)	-13.78	80.653	India (IN)	-3.73	49.413	Portugal(PT)	-3.15	44.574
	(0.00)	(0.00)		(0.01)	(0.01)		(0.03)	(0.02)
Barbados (BB)	-14.16	101.970	Ireland (IR)	-3.88	35.151	Peru (PE)	-5.164	100.670
	(0.00)	(0.00)		(0.00)	(0.14)		(0.99)	(0.00)
Belgium (BG)	-13.19	22.303	Israel (IS)	-6.45	98.350	Poland (PO)	-4.456	97.411
	(0.00)	(0.72)		(0.00)	(0.00)		(0.12)	(0.00)
Brazil (BR)	-5.77	77.564	Italy (IT)	-4.03	38.398	Spain (ES)	-9.77	68.606
	(0.45)	(0.00)		(0.00)	(0.07)		(0.00)	(0.00)
Canada (CA)	-6.03	81.477	Jordan (JO)	-8.35	62.969	Sweden (SD)	-8.10	77.826
	(0.00)	(0.00)		(0.00)	(0.00)		(0.00)	(0.00)
Columbia (CO)	-1.32	53.917	Japan (JP)	-13.80	99.320	Turkey (TK)	-7.115	49.511
	(0.69)	(0.00)		(0.00)	(0.00)		(0.00)	(0.01)
Denmark (DK)	-15.88	48.837	Korea (KO)	-10.17	180.340	Taiwan(TW)	-19.23	129.675
	(0.00)	(0.01)		(0.00)	(0.00)		(0.00)	(0.00)
Finland (FN)	-5.17	52.996	Luxembourg (LU)	-16.99	92.869	UK	-5.88	210.513
	(0.00)	(0.00)		(0.00)	(0.00)		(0.00)	(0.00)
France (FR)	-13.66	21.976	Mexico(MX)	-3.94	133.319	US	-10.18	193.921
	(0.00)	(0.74)		(0.02)	(0.00)		(0.00)	(0.00)
Germany (DE)	-6.50	52.890	Malaysia (MY)	-11.78	23.498	-	-	-
	(0.00)	(0.00)		(0.00)	(0.66)			
Greece (GR)	-3.62	56.790	Netherlands (NL)	-16.16	47.790	-	-	-
	(0.04)	(0.00)		(0.00)	(0.01)			

ADF test statistics for inflation rates and diagnostics for the multivariate ARCH-LM test. ADF test decisions are based on heteroscedasticity-robust  $p$ -values obtained by resampling as introduced by Cavaliere and Taylor (2008). ADF regressions include an intercept term and lag selection is according to the SIC. Multivariate ARCH-LM diagnostics refer to ARCH effects up to the third order in the vector residuals  $e_{it}$  obtained from model (2.2).

most sensibly associated with uncertainty. Such volatilities may be interpreted as representations of aggregate uncertainty. Grier and Perry (2000) point out that such expression of uncertainty accord with theoretical concepts of inflation or output uncertainty as discussed in contributions of Cukierman and Meltzer (1986) or Devereux (1989). For economies  $i = 1, \dots, N$ , the bivariate GARCH specification for conditional second moments in inflation and output is modeled as

$$H_{it} = C_i' C_i + A_i' e_{i,t-1} e_{i,t-1}' A_i + G_i' H_{i,t-1} G_i, \quad \text{with} \quad (2.1)$$

$$C_i = \begin{bmatrix} c_{i11} & c_{i12} \\ 0 & c_{i22} \end{bmatrix}, \quad A_i = \begin{bmatrix} a_{i11} & a_{i12} \\ a_{i21} & a_{i22} \end{bmatrix}, \quad G_i = \begin{bmatrix} g_{i11} & g_{i12} \\ g_{i21} & g_{i22} \end{bmatrix}, \quad H_{it} = \begin{bmatrix} h_{it}^{(\pi)} & h_{it}^{(\pi y)} \\ h_{it}^{(\pi y)} & h_{it}^{(y)} \end{bmatrix}.$$

Accordingly,  $H_{it}$  represents uncertainties in inflation and output, while elements of  $A_i$  ( $G_i$ ) characterize the impact of inflation and output growth surprises (past inflation

and output uncertainties) on  $H_{it}$ . The aforementioned inflation and output surprises, denoted as  $e_{it} = (e_{it}^{(\pi)}, e_{it}^{(y)})' = H_{it}^{1/2} u_{it}$ ,  $u_{it} \sim (0, I_2)$ , are implied by the VAR model

$$\begin{aligned} \begin{pmatrix} \pi_{it} \\ y_{it} \end{pmatrix} &= \mu + \sum_{p=1}^P \underbrace{\begin{bmatrix} \gamma_{ip}^{(\pi)} & \gamma_{ip}^{(y\pi)} \\ \gamma_{ip}^{(\pi y)} & \gamma_{ip}^{(y)} \end{bmatrix}}_{\Gamma_{ip}} \begin{pmatrix} \pi_{i,t-p} \\ y_{i,t-p} \end{pmatrix} + \sum_{q=1}^Q \underbrace{\begin{bmatrix} \psi_{iq}^{(oil)} & 0 \\ 0 & \psi_{iq}^{(\bar{y})} \end{bmatrix}}_{\Psi_{iq}} \begin{pmatrix} oil_{i,t-q} \\ \bar{y}_{t-q} \end{pmatrix} \\ &+ \underbrace{\begin{bmatrix} \lambda_i^{(\pi)} & \lambda_i^{(y\pi)} \\ \lambda_i^{(\pi y)} & \lambda_i^{(y)} \end{bmatrix}}_{\Lambda_i} \begin{pmatrix} \sqrt{h_{it}^{(\pi)}} \\ \sqrt{h_{it}^{(y)}} \end{pmatrix} + \begin{pmatrix} e_{it}^{(\pi)} \\ e_{it}^{(y)} \end{pmatrix}, \end{aligned} \quad (2.2)$$

where  $\mu$  denotes a vector of intercept terms. The parameter matrices  $\Gamma_{ip}$  captures linkages between the levels of inflation and output. It is often argued that, apart from idiosyncratic influences, these variables may be driven by factors that are determined from outside the domestic economy. Among the most important ones are oil prices or commonalities in business cycles among open economies (Ciccarelli and Mojon 2010). The latter determinant is expressed as  $\bar{y}_{t-1} = (1/N) \sum_{i=1}^N y_{i,t-1}$ , i.e. the average over economy-specific output growth rates. Potential impacts of  $oil_{i,t-1}$  and  $\bar{y}_{t-1}$  are represented by means of the parameters in  $\Psi_{iq}$ , whereas the effect of inflation and output uncertainties  $\sqrt{h_{it}^{(\pi)}}$  and  $\sqrt{h_{it}^{(y)}}$  is quantified in terms of the matrix  $\Lambda_i$ .

### 2.1.3 Implementation details

For all countries, the lag orders in (2.2) are chosen as  $P = 12$  and  $Q = 3$ . In (2.1), the  $2 \times 2$  parameter matrices  $A_i$  and  $G_i$  are fully parameterized, while its upper left elements  $a_{11}$  and  $g_{11}$  are restricted to be positive to guarantee the identification of parameters (Engle and Kroner 1995). Proceeding in this way we allow for cross-equation linkages of second order dynamics in (2.1). Formulating the bivariate GARCH model in (2.1) as a BEKK specification (Engle and Kroner 1995) obtains parameter estimates in terms of nonlinear functions. For this reason we report empirical results for estimates of  $A_i$  and  $B_i$  by reformulating BEKK-implied results in terms of the so-called half-vec form (Bollerslev, Engle and Wooldridge 1988) that enables straightforward interpretation.<sup>4</sup> The system as described in (2.2) and (2.1) is estimated for each

<sup>4</sup>For this purpose, define  $h_t \equiv vech(H_t)$  and  $\nu_t \equiv vech(e_t e_t')$ , omitting country indices for notational convenience. Then, equation (2.1) might be reformulated in terms of  $h_t = c + A^* \nu_{t-1} + G^* h_{t-1}$ , deriving model parameters  $M^* \in \{A^*, G^*\}$  from their counterparts  $M \in \{A, G\}$  in (2.1) as

country  $i = 1, \dots, 34$  individually by numerical optimization of the likelihood function regarding model disturbances  $u_{it}$ . Owing to potential nonnormality of  $u_{it}$ , resulting quantifications of causality coefficients are regarded as Quasi-Maximum-Likelihood estimates (Comte and Lieberman 2003).

### 2.1.4 Inference

To account for the heterogeneity inherent in country-specific inflation and output dynamics, the system (2.1) and (2.2) is estimated individually for each economy. To summarize the findings regarding parameters obtained in this way for particular economies, the mean-group (MG) estimation method as introduced by Pesaran and Smith (1995) is employed.

In the following, let  $\theta_i$  represent any parameter from (2.1) or (2.2) for economy  $i$ . The MG estimator then reads as

$$\bar{\theta} = \frac{1}{N} \sum_{i=1}^N \theta_i. \quad (2.3)$$

As an indication of whether relations in (2.1) and (2.2) feature qualitatively similar parameters over the cross section, we consider the following  $t$ -statistic:

$$t = \sqrt{N} \frac{\bar{\theta}}{\sigma}, \quad (2.4)$$

$$\text{where } \sigma = \sqrt{\frac{1}{N-1} \sum_{i=1}^N (\theta_i - \bar{\theta})^2}, \quad (2.5)$$

following Coakley, Fuertes and Smith (2001). According to Pesaran and Smith (1995), the estimator in (2.3) is suitable for dynamic models and data sets featuring relatively large cross-section and time dimensions. In contrast to most other panel techniques, moreover, it does not rely on a parameter homogeneity assumption. Furthermore, the cross sectional  $t$ -statistic in (2.4) allows for accurate inference even in cases of

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follows:

$$M^* = \begin{bmatrix} m_{11}^2 & 2m_{11}m_{21} & m_{21}^2 \\ m_{11}m_{12} & m_{21}m_{12} + m_{11}m_{22} & m_{21}m_{22} \\ m_{12}^2 & 2m_{12}m_{22} & m_{22}^2 \end{bmatrix}.$$

For example, noting that  $h_t = [h_{it}^{(\pi)} h_{it}^{(\pi y)} h_{it}^{(y)}]'$ , the lower left element given by  $m_{12}^2$ , where  $m_{12}^2 \in \{g_{12}^2, a_{12}^2\}$ , quantifies the nonnegative effect that lagged inflation uncertainty/lagged inflation shocks exerts on output uncertainty.

nonstationary variables and small to moderate degrees of correlation among economy-specific coefficient estimates (Coakley, Fuertes and Smith 2001, Ibragimov and Mueller 2010). Significance of (2.4) indicates that a particular effect holds in similar form across distinct types of economies. In the following, we will denote such effects as being "economically relevant".

## **2.2 Hypotheses about inflation and output dynamics**

Most of the potential relations between inflation, output and corresponding uncertainties in inflation and output are discussed in the theoretical literature. In the first part of this Section, we summarize theoretical hypotheses on causalities that we can empirically address in a world-wide context in the framework of the model described above. In the second part, we acknowledge that many contributions motivate causal effects for particular variables to arise as a matter of monetary policy. Accordingly, we summarize respective hypotheses stressing that particular effects might depend on the characteristics of respective economies.

The way in which linkages between inflation and output are understood has changed considerably since the initial formulation of the Phillips curve. Prominent contributions in this literature are Samuelson and Solow (1960) or Lucas and Sargent (1979), *inter alia*. While the existence of a long term tradeoff between the two quantities is highly controversial, most economists agree that a relation between inflation and measures of real output is present for the short- to medium term. This is documented in several empirical studies (e.g. King and Watson 1994, Galí and Gertler 1999, Sbordone 2005). Accordingly, short-run Phillips curve dynamics are an important building block in modern DSGE models (Smets and Wouters 2010). The relation between uncertainties and the conditional means of inflation and output has equivalently attracted considerable attention in the theoretical literature. The discussion on respective relations and their economic relevance, however, is still a matter of controversy. In Table 2.2, some of the most prominent hypotheses on the causal impact among the variables we consider are summarized. Not all of the enlisted effects are considered explicitly within our study.

Both the signs and directions of potential causal effects are in many cases controversial. Widely cited theoretical concepts of the relation between inflation and IU are Friedman (1977), Ball (1992), Cukierman and Meltzer (1986) or Devereux (1989). In Friedman (1977) and Ball (1992), it is argued that it is the level of inflation which gives rise to IU, whereas Cukierman and Meltzer (1986) and Devereux (1989) describe

Table 2.2: Theoretical hypotheses on causal effects and their dependence on  $\pi$

Hypothesis		Reference & Form		Effect in model (1)-(2)
$\pi$	$\overset{+}{\rightarrow} y$	IPC	BMR	$\sum_{p=1}^3 \gamma_{ip}^{(\pi y)}$ , $\sum_{p=1}^{12} \gamma_{ip}^{(\pi y)}$
$y$	$\overset{+}{\rightarrow} \pi$	PC	BMR	$\sum_{p=1}^3 \gamma_{ip}^{(y\pi)}$ , $\sum_{p=1}^{12} \gamma_{ip}^{(y\pi)}$
IU	$\overset{\pm}{\rightarrow} y$	FDS	G	$\lambda^{(\pi y)}$
IU	$\overset{\pm}{\rightarrow} \pi$	CMD	CMD	$\lambda^{(\pi)}$
OU	$\overset{\pm}{\rightarrow} y$	BP	-	$\lambda^{(y)}$
OU	$\overset{+}{\rightarrow} \pi$	DE	-	$\lambda^{(y\pi)}$
OU	$\overset{\pm}{\leftarrow} \text{IU}$	T	TF	$a_{21}^2, g_{21}^2, a_{12}^2, g_{12}^2$
$\pi$	$\overset{+}{\rightarrow} \text{IU}$	FB	-	implicit

Hypotheses on inflation and output dynamics. In the column 'Reference & Form', references for theoretical hypotheses on causalities (left part) and their relation to monetary performance (right part) are given in abbreviated form. Namely, '(I)PC' is short for (Inverted) Phillips curve, while 'BMR' is Ball, Mankiw, and Romer (1988). Furthermore, 'FDS' is Friedman (1977) or Dotsey and Sartre (2000), 'CMD' is Cukierman and Meltzer (1986), Devereux (1989), 'T' is short for Taylor (1994), 'DE' is Devereux (1989), 'TF' is Taylor (1994) and Fuhrer (1997), while 'FB' is Friedman (1977) and Ball (1992). For notational convenience, we will denote the 'TF' implied second order effects as  $a^{(y\pi)}$ ,  $g^{(y\pi)}$ ,  $a^{(\pi y)}$ ,  $g^{(\pi y)}$  later on.

scenarios where causality points in the opposite direction. Positive output effects of IU are described by e.g. Dotsey and Sarte (2000), who introduce a model in which IU leads to increases in investment through precautionary savings. In contrast, Friedman (1977), for example, describes a negative effect of IU on output. He argues that inflation uncertainty tends to reduce the effectiveness of the price mechanism in allocating resources efficiently. Moreover, Sargent (1993) postulates that a lack of central bank credibility is detrimental to output, where credibility might be associated with low levels of IU (Giordani and Söderlind 2003). With respect to the relation between output and OU, Black (1987) postulates that higher output uncertainty increases output growth since the latter might reflect an engagement in riskier projects. In contrast, according to the model of Pindyck (1991), OU has a negative influence on output growth since irreversible investment projects tend to be postponed under higher uncertainty. In Devereux (1989), output uncertainty increases the average rate of inflation by an implicit reduction of wage indexation which provides an incentive for policymakers to generate inflation surprises more frequently.

When hypotheses like the ones stated above are investigated, it is usually implicitly assumed that the relation in question holds invariably for distinct economies. The way in which many assertions regarding causality are formulated, however, suggests that



some of them might only prevail in some economies that share certain characteristics, in particular regarding the implementation and the success of monetary policy. For instance, as the average level of inflation influences the price adjustment frequency, it might affect the output-inflation trade-off (Ball, Mankiw, and Romer 1988). Further examples include the widely cited models of Cukierman and Meltzer (1986) or Devereux (1989), where only IU in excess of a certain level might gain potential to impact on inflation. Other such examples are Gillman et al. (2004) who discuss a potentially inflation-dependent effect of IU on output and also provide empirical evidence for their conjecture. Fountas et al. (2004) and Neanidis and Savva (2011) document that the constitutional frame of monetary policy can have an influence on empirical estimates of a causal impact of IU on growth. Moreover, Taylor (1994) and Fuhrer (1997) assert that the relation between IU and OU might depend on the level of inflation. In particular, Fuhrer (1997) points out that a tradeoff between IU and OU is particularly relevant for low-inflation economies. Gylfason and Herbertsson (2001) postulate that in low inflation countries, growth is more sensitive to inflation changes than in high inflation countries.

Accordingly, though the literature does not provide consensus on both the sign and the direction of potential causalities, it appears rather unanimous in stressing the potential role of monetary performance as an important source of systematic cross sectional heterogeneity. Hence we focus on two empirical questions. On the one hand, we address if particular hypotheses appear economically relevant across the full cross section. On the other hand, we examine how sensitive associated effects are with respect to the characteristics of economies. For this purpose we consider average causal relations in subsamples of the cross section which are characterized by distinct levels of inflation. We argue that the overall level of inflation over several years should be associated with the reliability of the respective economy's monetary authority and its success in anchoring inflation expectations.

## **2.3 Results**

### **2.3.1 Causal linkages**

In this Section, we analyze the outcomes of the model (2.1)-(2.2) over the considered set of economies. By means of studying parameter variation across economies, we investigate the strength of evidence for the hypotheses summarized in Table 2.2. Furthermore, we discuss mean impulse response functions (IRFs) regarding surprises in

inflation and output.<sup>5</sup> Proceeding in this way we highlight potential causal linkages with emphasis on a more detailed description of the shock response of the variables in question. Mean group estimates regarding relations between output and inflation from VAR estimates in (2.2) are reported in Table 2.3. To summarize the informational content of AR coefficients, we further provide cross sectional averages of accumulated country specific impacts obtained over the initial three (twelve) lags. The Table also contains the effects of uncertainty variables in  $H_{it}$  on the levels of output and inflation, whereas accumulated IRFs are depicted in Figure 2.1. A brief discussion of diagnostic tests regarding the model specification concludes.

The results given in Table 2.3 show that inflation and output are significantly interrelated. In case of the impact of productivity growth on inflation quantified by  $\bar{\gamma}_{ip}^{y\pi}$ , most lag coefficients are positive, where four out of 12 are significant. The lag parameters  $\bar{\gamma}_{ip}^{\pi y}$ , which capture the effect of inflation on output tend to be negative, while being significant in two cases. Cross sectional averages of accumulated quarterly impacts are  $\bar{\gamma}_3^{(y\pi)} = 0.016$  and  $\bar{\gamma}_3^{(\pi y)} = 0.25$ , respectively, where only the former estimate is significant. Accumulating over all lag coefficients up to one year, however, results in estimates  $\bar{\gamma}_{12}^{(y\pi)} = 0.04$  and  $\bar{\gamma}_{12}^{(\pi y)} = -0.96$  that are both significant at the 1% level. Thus, a 1% increase in output growth has an associated inflation effect of only 4 basis points, whereas an increase in inflation by 1% is followed by a decrease in output of almost 1%.<sup>6</sup> In terms of empirical magnitude, therefore, the estimated effect of inflation on output appears by far more relevant than the impact that output exerts on inflation. However, the influence that output exerts on inflation is found to be more systematic across economies and also takes, on average, less time to become evident than the reverse effect. This impression is confirmed by the accumulated (long run) IRFs depicted in Figure 2.1, where a surprise in output tends to be followed by a rise in the price level pretty quickly. Though occurring with a time lag, however, a large fraction of the impact of a surprise in inflation on output materializes during the first year after the shock.

Our results might imply that, though the effect of output on inflation can be found almost immediately and reliably across economies, the consequences of excess inflation might be of greater economic importance in modern economies.

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<sup>5</sup>The underlying country specific impulse response functions for GARCH-M VAR models are calculated according to Elder (2004).

<sup>6</sup>A negative impact of inflation on output growth is in line with empirical studies such as Barro (1996), Temple (2000), Khan (2001) or Gillman and Kejak (2005).

Table 2.3: Mean group estimation of equation (1)

Eq	$\pi_t$			$y_t$		
	$y_{t-l}$	$\pi_{t-l}$	$oil_{t-l}$	$y_{t-l}$	$\pi_{t-l}$	$\bar{y}_{t-l}$
1	0.0051 (1.366)	0.1806 (4.196)	0.0047 (7.775)	-0.3815 (-7.235)	0.1572 (1.138)	0.1855 (2.982)
2	0.0042 (1.105)	0.0238 (1.233)	0.0003 (0.000)	-0.2249 (-5.886)	-0.0392 (-0.416)	0.3636 (0.364)
3	0.0064 (3.047)	0.0644 (4.067)	0.0006 (1.050)	-0.1350 (-4.449)	0.1330 (1.087)	0.3791 (5.920)
4	0.0107 (4.504)	0.0416 (3.493)	-	-0.0447 (-1.938)	0.0177 (0.146)	-
5	0.0047 (1.316)	0.0402 (3.203)	-	0.0205 (1.294)	-0.2833 (-2.322)	-
6	0.0041 (2.019)	0.0606 (4.706)	-	0.0457 (2.214)	-0.1228 (-1.284)	-
7	0.0026 (1.070)	0.0465 (3.463)	-	-0.0015 (-0.079)	-0.1474 (-1.421)	-
8	0.0061 (2.151)	0.0526 (3.841)	-	-0.0079 (-0.417)	-0.0610 (-0.579)	-
9	-0.0034 (-1.213)	0.0400 (3.205)	-	0.0351 (1.831)	-0.1324 (-1.047)	-
10	-0.0006 (-0.237)	0.0585 (3.533)	-	-0.0296 (-1.950)	-0.1446 (-1.033)	-
11	-0.0017 (-0.692)	0.0647 (4.629)	-	-0.0403 (-2.325)	-0.0046 (-0.040)	-
12	-0.0001 (-0.039)	-0.0725 (-2.956)	-	-0.0273 (-1.484)	-0.3370 (-2.375)	-
$\sum_{l=1}^3$	0.0157 (2.170)	0.2688 (5.317)	0.0056 (5.507)	-0.7415 (-7.604)	0.2509 (1.206)	0.9282 (6.476)
$\sum_{l=1}^{12}$	0.0381 (3.560)	0.6010 (11.966)	-	-0.7916 (-6.128)	-0.9646 (-3.424)	-
$h_t$	0.0329 (0.647)	-0.4616 (-1.298)	-	-0.2146 (-0.507)	-2.9505 (-2.848)	-

Mean group VAR(12) diagnostics for inflation and output growth summarizing the informational content of individual point estimates of VAR(12) models with BEKK(1,1)-GARCH-M type residuals estimated for 34 economies.  $t$ -statistics according to (2.4) are given in parentheses.  $\sum_{l=1}^3$  ( $\sum_{l=1}^{12}$ ) is short for lag coefficients accumulated up to 3 (12) months.

Hence, one might conjecture that, in modern economies that tend to be characterized by well anchored inflation expectations, inflation is not overly sensitive to economic activity, but output growth tends to be quite vulnerable to inflation. The estimated inflation effect on output appears even more relevant when it is compared, for instance, to the accumulated effect of the global average output growth on domestic growth rates  $\bar{\psi}_3^{(\bar{y})}$ . The response of domestic growth to the world business cycle

has the same magnitude as the respective inflation effect. Hence, on average over the considered cross section, the output loss due to a moderate increase in inflation might be comparable to the one incurred in case of a 1% decrease in average world output growth.

Moreover, we find considerable support for an adverse effect of IU on output growth. In 27 out of 34 economies, we observe negative coefficient estimates for the linkage running from IU to output growth rates. The average effect of IU on output growth is  $\bar{\lambda}^{(\pi y)} = -2.95$ , with the associated  $t$ -ratio indicating significance at the 1% level. To provide a benchmark, annualized US inflation and production have unconditional means of about 2.7 % and 1.7%, whereas the unconditional standard deviations amount to 0.27 % and 0.8% for inflation and output, respectively. Accordingly, the statistically significant negative impact of inflation and inflation uncertainty on industrial production also appear to be relevant in terms of economic magnitude. Put differently, an effect of both inflation and IU on the real economy can be observed for a considerable part of the cross section. In the related literature, the effects of inflation and IU are often not treated separately as it is often presumed that a strong linkage exists between the two variables. The results in Table 2.3, however, suggest that potential disadvantages may arise from IU in addition to effects associated with the level of inflation.

Table 2.4: Mean group estimates for equation (2)

half-vec dynamics of VAR(12) residuals						
$Eq$	$e_{t-1}^{(\pi)2}$	$e_{t-1}^{(\pi)} e_{t-1}^{(y)}$	$e_{t-1}^{(y)2}$	$h_{t-1}^{(\pi)}$	$h_{t-1}^{(\pi y)}$	$h_{t-1}^{(y)}$
$h_t^{(\pi)}$	0.3173 (7.485)	0.0106 (2.154)	0.0014 (2.107)	0.3422 (6.049)	0.0050 (1.333)	0.0004 (2.529)
$h_t^{(\pi y)}$	0.0046 (1.918)	0.1796 (8.625)	-0.0000 (-0.009)	0.0049 (0.451)	0.3436 (7.365)	0.0027 (1.282)
$h_t^{(y)}$	0.0023 (2.814)	0.0104 (1.873)	0.1697 (6.417)	0.0058 (1.703)	0.0143 (0.795)	0.4376 (8.358)

Mean group BEKK implied half-vec diagnostics for inflation and output uncertainties summarizing the informational content of individual point estimates of (2.1) for 34 economies.  $t$ -statistics according to (2.4) are given in parentheses.

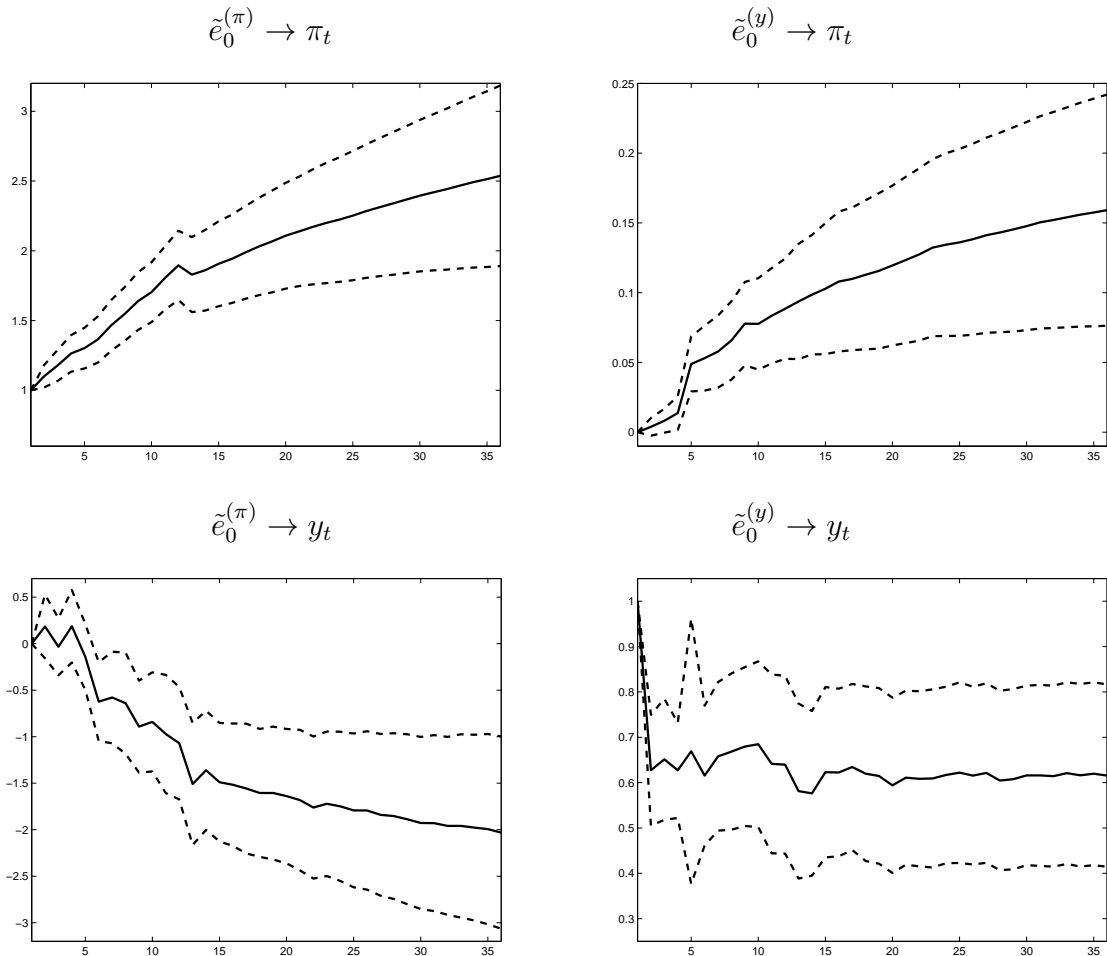


Figure 2.1: Average impulse response functions (accumulated)

Mean group accumulated impulse response functions obtained by averaging country specific impulse response function estimates. Country specific impulse response function for GARCH-M VAR models are calculated according to Elder (2004). Confidence bands are calculated as 2 times the cross sectional standard deviation divided by  $\sqrt{N}$ .

On the other hand, the hypotheses which suggest an impact of IU on inflation cannot be confirmed noting that the coefficient average  $\bar{\lambda}^{(\pi)}$  is insignificant even at the 10% nominal level. Moreover, in contrast to Fountas, Karanasos and Kim (2002), we do not find evidence for the hypothesis that OU is negatively related to output growth, i.e. stabilization objectives do not seem to be beneficial regarding the level of output growth.

Summarizing evidence regarding linkages among IU and OU in an analogous way, corresponding mean group parameters are reported in Table 2.4. Due to the positivity restriction imposed on the spillover parameters in the framework of the BEKK

specification, significance of these coefficients is less meaningful from the perspective of economic relevance. In contrast, the insignificance of the effect of IU on OU suggest rather strong evidence against the hypothesis that IU impacts on OU, which is documented by Grier et al. (2004). In line with Grier et al. (2004), however, the reverse effect of OU on IU and lagged idiosyncratic influences are throughout significant, albeit numerically small.

Finally, we document that modeling relations between the nominal and real economy by means of the particular model (2.1)-(2.2) is not only guided by the demand for a parametrization of the economic hypotheses in question. It seems also justified from an empirical perspective, i.e. by means of diagnostic tests regarding the model specification. Firstly, multivariate ARCH-LM tests are considered to judge the effectiveness of the specification (2.1) in expressing the evolution of inflation and output uncertainties. Second, we employ multivariate Breusch-Pagan LM test statistics to test for serial correlation up to order 6 in the standardized residuals  $u_{it} = H_{it}^{-1/2} e_{it}$ . For almost all countries in the cross section, neither ARCH effects nor autocorrelations can be detected in  $u_{it}$ . Table 2.5 summarizes the diagnostic test results across economies.

Table 2.5: Standardized Residuals: Multivariate diagnostics

	SC	ARCH		SC	ARCH		SC	ARCH
AG	19.1	65.7	GR	47.1*	54.3	NL	26.6	61.7
AT	26.1	80.7	HN	20.7	34.1	NW	26.3	31.9
BB	28.6	73.5	IN	24.2	52.4	PT	33.3	39.0
BG	39.8	59.8	IR	28.6	33.6	PE	40.0	73.8
BR	36.4	63.40	IS	56.2*	62.8	Po	21.2	57.4
CA	48.1*	31.9	IT	41.1	43.4	ES	30.0	48.9
CH	18.0	64.7	JO	20.1	54.9	SD	35.3	30.0
CO	29.7	48.0	JP	27.5	30.6	TK	24.8	56.8
DK	31.8	47.3	KO	23.8	64.9	TW	38.0	47.0
FN	38.8	47.6	LU	38.3	79.2	UK	43.9*	100.1*
FR	35.7	43.9	MX	24.4	27.1	US	31.9	45.0
DE	27.8	36.6	MY	38.2	43.9	-	-	-

Multivariate residual diagnostics for country specific system estimates. SC (ARCH) is short for the multivariate Breusch-Pagan LM-test statistic for serial correlation (the LM-test statistic for multivariate ARCH effects) up to order 6 in the standardized vector residuals  $\hat{u}_{it}$ . Significance at the 1% level is indicated by asterisks.

### 2.3.2 Causality analysis for distinct inflation environments

In addition to the discussion of results for the entire cross section, we investigate results obtained by consideration of particular subgroups in the following.

Several theoretical contributions postulate that causal relationships among inflation, output and respective uncertainties might be influenced by certain characteristics of the economies in question. In particular, one might consider the average magnitude of the inflation rate over some longer time period as a determinant of causal linkages. The average inflation rate thereby serves as an indication for the reliability of the respective economy's monetary authority and its success in anchoring inflation expectations. To investigate on whether estimates differ systematically across subgroups of economies, we examine coefficient averages regarding particular subgroups of economies. Respective subgroups are defined with respect to their members' average inflation rates,  $\bar{\pi}_i^{(*)} = (1/T) \sum_{t=1}^T \pi_{it}$ , where  $*$   $\in$   $\{l, m, h\}$  indicates membership in either group of low, medium or high inflation economies. The latter is characterized by

$$\bar{\pi}_i^{(l)} \leq b_l < \bar{\pi}_i^{(m)} \leq b_m < \bar{\pi}_i^{(h)}, \quad (2.6)$$

with  $b_l = 2.5$ ,  $b_m = 5$  as boundary values. Subgroup specific parameter averages  $\bar{\theta}_*$  obtain as averages over a number of  $N_*$  economies as in the case of (2.4) for the entire cross section. Moreover,  $\sigma_*$  is obtained for respective cross section members in analogy to (2.5). To assess the significance of differences in coefficients in a pairwise subsample comparison we employ the two-sample  $t$ -statistic

$$t_{a,b} = \frac{\bar{\theta}_a - \bar{\theta}_b}{\sqrt{\frac{1}{N_a-1}\sigma_a^2 + \frac{1}{N_b-1}\sigma_b^2}}, \text{ where } \{a, b\} \in \{l, m; l, h; m, h\}. \quad (2.7)$$

Table 2.6 provides mean parameter estimates for Model (1)-(2) for the latter subgroups along with respective pairwise two-sample  $t$ -statistics. Most of the effects documented in the previous section cannot be rejected to be equal across economies with distinct inflation experiences. However, in low inflation economies, real effects of inflation appear to be stronger; Compared to medium and high inflation economies, inflation tends to be followed by higher economic growth in the subsequent quarter. In turn, over the rest of the year, the reduction of output growth is significantly more pronounced both in terms of size and significance of respective coefficients.

Table 2.6: subsample estimates and mean differences

eq	coef	$l$	$m$	$h$	$l$ vs $m$	$l$ vs $h$	$m$ vs $h$
$\pi$	$\sum_1^3 \gamma^{(y\pi)}$	0.0202 (2.149)	0.0139 (1.199)	0.1951 (1.147)	0.0063 (0.434)	-0.1749 (-1.145)	-0.1811 (-1.037)
	$\sum_1^{12} \gamma^{(y\pi)}$	0.0368 (2.147)	0.0387 (2.324)	0.1630 (1.371)	-0.0019 (-0.080)	-0.1262 (-1.167)	-0.1243 (-1.012)
	$\sum_1^3 \gamma^{(\pi)}$	0.0404 (0.677)	0.2636 (3.840)	0.4981 (6.645)	-0.2232 (-2.516)	-0.4577 (-4.948)	-0.2345 (-2.351)
	$\sum_1^{12} \gamma^{(\pi)}$	0.4132 (7.786)	0.5475 (5.640)	0.8471 (19.593)	-0.1343 (-1.320)	-0.4340 (-6.326)	-0.2996 (-2.987)
	$\lambda^{(y\pi)}$	0.0082 (0.153)	0.0197 (0.133)	0.1478 (1.227)	-0.0115 (-0.082)	-0.1395 (-1.142)	-0.1280 (-0.694)
	$\lambda^{(\pi)}$	0.0557 (0.248)	-1.2147 (-1.056)	-0.2601 (-1.141)	1.2703 (1.257)	0.3158 (1.003)	-0.9546 (-0.875)
	$\sum_1^3 \delta^{(oil)}$	0.0044 (3.563)	0.0083 (3.268)	0.0036 (2.640)	-0.0039 (-1.517)	0.0008 (0.458)	0.0047 (1.724)
	$y$	$\sum_1^3 \gamma^{(y)}$	-0.7178 (-4.054)	-0.7196 (-4.435)	-0.6191 (-2.817)	0.0018 (0.007)	-0.0986 (-0.361)
$\sum_1^{12} \gamma^{(y)}$		-0.7446 (-3.423)	-0.6442 (-2.788)	-0.7403 (-2.780)	-0.1004 (-0.321)	-0.0043 (-0.013)	0.0961 (0.277)
$\sum_1^3 \gamma^{(\pi y)}$		0.8861 (2.045)	-0.0857 (-0.268)	-0.1442 (-1.260)	0.9718 (1.746)	1.0303 (2.176)	0.0585 (0.183)
$\sum_1^{12} \gamma^{(\pi y)}$		-1.8931 (-3.363)	-0.7182 (-1.863)	-0.0824 (-0.352)	-1.1748 (-1.653)	-1.8106 (-2.846)	-0.6358 (-1.477)
$\lambda^{(y)}$		0.1973 (0.522)	-0.8223 (-0.656)	-0.2569 (-0.479)	1.0196 (0.888)	0.4542 (0.724)	-0.5654 (-0.440)
$\lambda^{(\pi y)}$		-2.2340 (-1.980)	-5.5277 (-1.905)	-1.4148 (-1.206)	3.2937 (1.188)	-0.8192 (-0.513)	-4.1129 (-1.396)
$\sum_1^3 \delta^{(\bar{y})}$		1.0705 (4.975)	0.7117 (2.890)	0.7040 (2.867)	0.3588 (1.124)	0.3665 (1.152)	0.0077 (0.023)
$h^\pi$		$a^{(y\pi)}$	0.0011 (1.714)	0.0031 (1.511)	0.0001 (2.107)	-0.0020 (-1.087)	0.0009 (1.426)
	$a^{(\pi)}$	0.3252 (4.713)	0.1983 (2.973)	0.3904 (5.580)	0.1269 (1.324)	-0.0652 (-0.675)	-0.1921 (-2.029)
	$g^{(y\pi)}$	0.0001 (2.560)	0.0011 (2.055)	0.0002 (2.049)	-0.0009 (-2.082)	-0.0000 (-0.109)	0.0009 (1.887)
	$g^{(\pi)}$	0.2498 (2.873)	0.4229 (4.057)	0.3782 (3.557)	-0.1731 (-1.315)	-0.1285 (-0.967)	0.0446 (0.306)
	$h^y$	$a^{(y)}$	0.1684 (3.099)	0.1699 (3.684)	0.2204 (3.751)	-0.0015 (-0.021)	-0.0520 (-0.664)
$a^{(\pi y)}$		0.0014 (1.406)	0.0037 (1.771)	0.0019 (1.678)	-0.0023 (-1.074)	-0.0005 (-0.319)	0.0018 (0.793)
$g^{(y)}$		0.3560 (3.746)	0.5219 (5.980)	0.4111 (4.618)	-0.1659 (-1.278)	-0.0550 (-0.426)	0.1108 (0.909)
$g^{(\pi y)}$		0.0013 (1.502)	0.0121 (1.200)	0.0053 (1.015)	-0.0108 (-1.251)	-0.0040 (-0.840)	0.0068 (0.628)

Mean group diagnostics for model (2.1)-(2.2) obtained for distinct subgroups of low ( $l$ ), medium ( $m$ ) and high ( $h$ ) inflation economies defined according to (2.6) (left hand side). The right hand side provides two-sample  $t$ -test statistics for pairwise comparisons of coefficient subgroup means according to (2.7).



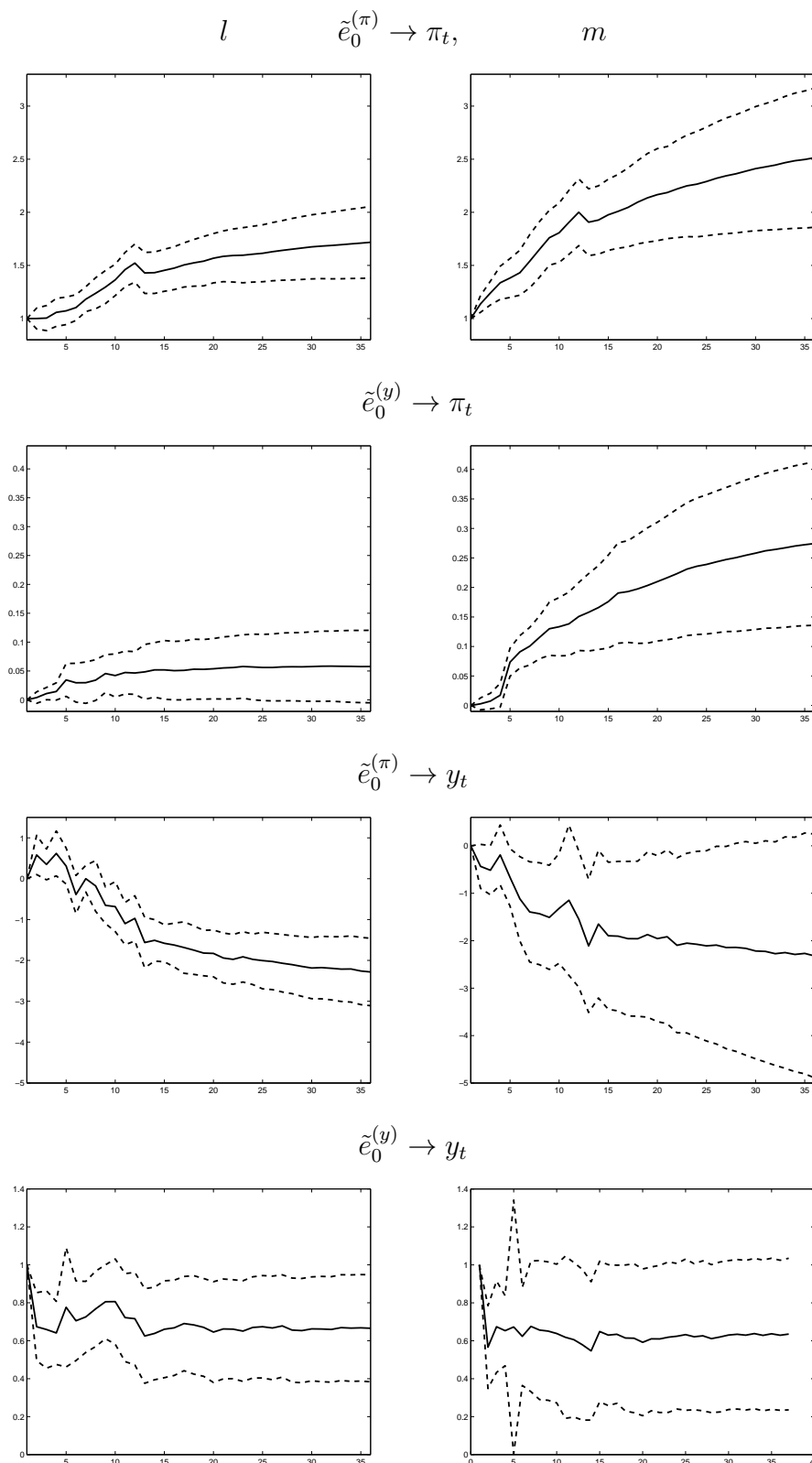


Figure 2.2: Average IRFs for low and medium inflation economies (accumulated)

Mean group accumulated IRFs for economies with average inflation below 2.5% (between 2.5% and 5%) in the left (right) panel. Country specific IRFs for GARCH-M VAR models are derived according to Elder (2004). Confidence bands are 2 times the group specific cross sectional standard deviation divided by  $\sqrt{K}$ ,  $K = \{13; 10\}$ .

This might be a result of scale effects, i.e. in economies with strongly accelerating prices, the real impact of an additional percentage rise in inflation is comparably smaller than an increase in an economy which has been characterized by low average inflation beforehand. Impulse responses depicted in Figure 2.2 illustrate the characteristic dynamics of output that tend to follow inflation shocks. Comparing the IRFs for low- and intermediate inflation regimes<sup>7</sup> it can be seen that the relation is most clear-cut for the low inflation economies, which exhibit lower standard deviations among their subgroup. In turn, IRFs further suggest that inflation tends to rise more markedly after a positive output shock in medium inflation economies compared to low inflation economies, which might be due to imperfect anchoring of inflation expectations.

Interestingly, the effect of IU on output growth does not significantly differ across economies with distinct inflation experiences. Influences of predetermined variables  $oil_{t-1}$  and  $\bar{y}_{t-1}$  are throughout unaffected with respect to the average inflation level. Moreover, volatility spillovers as obtained in the previous section also appear to be unaffected by the average level of inflation in the respective economies. Both the insignificance of the IU impact on output growth uncertainty and the significance of the other coefficients expressing uncertainty linkages are robust with respect to the consideration of distinct subgroups.

We conclude that there is only moderate evidence that particular interrelationships of those enlisted in Table (2.2) are dominated by issues of monetary stability. Moreover, effects that appear to depend on monetary stability such as the impact of inflation on output growth only differ in terms of magnitudes rather than in terms of sign or significance.

## 2.4 Conclusion

In this chapter, we assess common patterns in the interaction of inflation, output and their uncertainties. We focus on the recent times when inflation targeting has been the predominantly adopted rule of monetary policy and consider a large cross section of 34 developed and emerging economies. Considering economy-specific estimates obtained from standard bivariate GARCH-in-mean models, we infer on the economic significance of causality hypotheses from the theoretical literature from a cross sectional perspective. We document marked influences of nominal quantities

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<sup>7</sup>We do not consider impulse response functions for higher inflation economies. Given that inflation rates in 6 out of 11 high inflation economies are non-stationary, an aggregation of country specific impulse responses appears hardly meaningful.

on output growth. Both the level of inflation and the associated uncertainty tend to lower output growth across a wide range of distinct economies. We further find that theoretical arguments regarding the dependency of causal linkages on the monetary policy framework are relevant for some of the relations we investigate. Particularly, the effects of inflation on output appear most pronounced for economies with low average inflation rates. However, the negative impact of inflation uncertainty on output growth appears to hold irrespective of the monetary regime. Furthermore, there exist significant spillover effects among uncertainties in output and inflation. The effect, however, seems to point from output uncertainty to inflation uncertainty, whereas the opposite impact is not significant.

To summarize, we find lower and stable inflation rates to be beneficial for output growth. This relation appears to hold irrespective of whether developed or economies in transition are considered.

# Chapter 3

## Capital market integration

### 3.1 Motivation

Capital market integration is a core issue of international economics since it improves the allocation of investment (Laeven 2003, Galindo, Schiantarelli and Weiss 2007, Abiad, Oomes and Ueda 2007), lowers the volatility of consumption (Bekaert, Harvey and Lundblad 2006) and fosters growth (Rousseau 2003, Bekaert, Harvey and Lundblad 2005, Bekaert, Harvey, Lundblad and Siegel 2007).<sup>1</sup> In the literature, several approaches have been applied to assess capital market integration. These are based on the international correlation of consumption (Obstfeld 1994, Bekaert, Harvey and Lundblad 2006), the relation of savings and investment (Feldstein and Horioka 1980, Blanchard and Giavazzi 2002), international capital flows (Bekaert, Harvey, and Lumsdaine 2003, Baele, Ferrando, Hördahl, Krylova and Monnet 2004) or the law of one price (e.g. Frankel 1992, Bekaert and Harvey 1995, Baele et al. 2004, Hardouvelis, Malliaropulos and Priestley 2006). The most direct approaches to measure capital market integration are based on the law of one price, which implies that capital can be raised at equal global cost. Focussing on the cost of external capital, studies in the field of international macroeconomics have traditionally employed interest rate parity conditions in covered, uncovered or real terms to test for capital market integration (Frankel 1992).<sup>2</sup>

In this chapter, we test for capital market integration in terms of the real interest rate parity (RIP). Under capital market integration, free flow of capital and technology diffusion tend to equalize the cost of long term capital and the marginal return

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<sup>1</sup>See also Prasad, Rogoff, Wei and Kose (2003) for negative evidence on growth effects.

<sup>2</sup>Another important approach based on the law of one price is followed in international finance, where asset pricing models are used to uncover if the cost of equity capital is determined with respect to global or local risks (see e.g. Bekaert and Harvey 1995, Hardouvelis et al. 2006, Carrieri, Errunza and Hogan 2009).

on national production possibilities (Obstfeld and Taylor 2002). Accordingly, the RIP indicates the efficiency of capital allocation and the exploitation of comparative advantages. The RIP further suggests that goods markets are integrated since it jointly implies the uncovered interest rate parity and the ex-ante relative purchasing power parity (UIP and PPP, hereafter).

In light of liberalized capital markets the weak evidence for the RIP has puzzled economists. Initially, Mishkin (1984), Cumby and Mishkin (1986) and Merrik and Saunders (1986) rejected the equality of international real interest rates (RIRs). Subsequently, research focused on the long run convergence of international RIRs (Edison and Pauls 1993, Goodwin and Grennes 1994, Moosa and Bhatti 1995, Chinn and Frankel 1995, Mac Donald and Nagayasu 2000, Obstfeld and Taylor 2002, Goldberg 2003, Mancuso, Goodwin and Grennes 2003, Ferrera and Ledesma 2007). While RIRs are often found to be cointegrated, long run parity relationships hardly receive empirical support. Likewise, cross country RIR differentials (RIDs) are often diagnosed to be non-stationary.

The empirical literature still suffers from three caveats. Firstly, with the exceptions of Fujii and Chinn (2001) and Obstfeld and Taylor (2002), empirical contributions rely on short maturity RIRs, typically 3 months. For inference on capital market integration, however, long term RIRs should be considered since they most directly reflect financing costs for capital investments (Obstfeld and Taylor 2002). Furthermore, the use of long term instruments reduces the potential shortcomings invoked by the failure of the ex-ante relative PPP and the UIP in the short run (Fujii and Chinn 2001, Obstfeld and Taylor 2002).<sup>3</sup> The relative PPP is accepted as a reasonable assumption only over long time horizons (Taylor and Taylor 2004). Similarly, the UIP receives stronger support for long term interest rates (Chinn and Meredith 2004). Hence, it is more likely that both conditions hold ex-ante if one considers financing of long term capital investments. Secondly, the studies to investigate the convergence of long term RIRs tend to suffer from imprecise RIR approximations.<sup>4</sup> RIRs are throughout determined by means of the Fisher hypothesis (Fisher 1930) which requires an approximation of expected inflation and the absence of inflation risk. To measure expected inflation, studies proceed from the assumption of rational expectations (e.g. Fujii and Chinn 2001 or Obstfeld and Taylor 2002) or employ model implied expectations (e.g. Al-Awad and Goodwin 1998, Fujii and Chinn 2001). Especially for ex-ante long run inflation, however, these approaches fail to guarantee

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<sup>3</sup>Empirical findings in Chung and Crowder (2004) attribute the failure of the RIP to the violation of both conditions.

<sup>4</sup>Notably, any systematic bias implies nonzero RIDs in presence of market integration.

idiosyncratic approximation errors at the monthly or quarterly frequency.<sup>5</sup> Moreover, the disregard of inflation risk is another source of systematic measurement errors. Thirdly, with the exceptions of Chinn and Frankel (1995) and Goodwin and Grennes (1994), studies rarely investigate market integration explicitly from the perspective of higher dimensional systems. The system approach provides one further testable rank condition that is necessary (though not sufficient) for capital market integration.

We assess capital market integration for a subsystem of the G7 by testing for the convergence of *observed* ex-ante long term RIRs as implied by treasury inflation protected securities (TIPS). The sample covers monthly RIR quotes for Canada (CA), France (FR), the UK and the US over the period 2000/1-2010/1. At first, we apply unit root tests to RIDs. In addition, we test for cointegration among RIRs and provide LR statistics for overidentifying parameter restrictions implied by the RIP. Our results suggest that CA, FR, UK and US capital markets are integrated. While all RIDs are diagnosed stationary around zero, bivariate cointegration analysis also supports the long run RIP. Full dimensional cointegration analysis suggests global linkage to the US market. Moreover, we cannot reject that the estimated cointegration relationships reflect the integration of CA, FR, UK and US capital markets. Addressing the lack of empirical support that the RIP receives in related studies, we reconsider our analysis by employing conventional RIR approximations. We find evidence that both the failure of short run ex-ante relative PPP or UIP, and presumably weak approximations of long term RIRs might induce an empirical rejection of the RIP in presence of capital market integration.

In Section 3.2 we briefly discuss the economic theory, the hypotheses and the reasoning behind the test specifications. Empirical results are provided in Section 3.3. In Section 3.4 capital market integration is reinvestigated in terms of conventional RIR measures. Section 3.5 summarizes and concludes.

## 3.2 Testing for capital market integration

### 3.2.1 On RIP and inflation indexed interest rates

In this study, capital market integration is diagnosed with reference to the law of one price. Two capital markets are integrated if the expected real cost of capital and, implicitly, the expected marginal yields on investment tend to equalize in the long run. This holds under free flow of capital and frictionless markets if  $i$ ) assets denoted

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<sup>5</sup>In particular, ex-post ten year inflation might not provide an unbiased measure of ex-ante ten year inflation.

in distinct currencies are regarded as substitutes and *ii*) goods market integration leads to expected exchange rate changes that reflect expected inflation differentials.

Consider the UIP condition,<sup>6</sup>

$$\xi_{t,k}^{(i)} - \xi_{t,k}^{(j)} = E_t \left[ s_{t+k}^{(i,j)} - s_t^{(i,j)} \right], \quad (3.1)$$

where  $\xi_{t,k}^{(\bullet)}$  is the  $k$ -period interest rate in country  $\bullet = i, j$ , and  $s_{t+k}^{(i,j)} - s_t^{(i,j)}$  is the  $k$ -period change of the log spot exchange rate, i.e. the price of currency  $i$  per units of currency  $j$ . The expectation operator  $E_t$  conditions on all information available in time  $t$ . Furthermore, let  $\pi_{t,k}^{(\bullet)}$  denote the inflation rate in country  $\bullet = i, j$  prevailing from period  $t$  to  $t + k$  and define the ex-ante RIR as  $E_t r_{t,k}^{(\bullet)} = \xi_{t,k}^{(\bullet)} - E_t \pi_{t,k}^{(\bullet)}$  (Fisher 1930). By virtue of ex-ante relative PPP the right hand side expectation in (3.1) obtains as  $E_t \left[ s_{t+k}^{(i,j)} - s_t^{(i,j)} \right] = E_t \pi_{t,k}^{(i)} - E_t \pi_{t,k}^{(j)}$ . Hence, UIP and ex-ante relative PPP jointly imply the equality of ex-ante RIRs, i.e.

$$E_t r_{t,k}^{(i)} = E_t r_{t,k}^{(j)}. \quad (3.2)$$

In related studies, RIRs are obtained by subtracting ex-ante inflation from nominal interest rates. Accordingly, RIR approximations are implicitly subject to two additional sources of uncertainty, namely the determination of expectations and the disregard of inflation risk premia. In this study, ex-ante RIRs  $r_{t,k}^{(\bullet)}$ ,  $\bullet = i, j$  are market quotes of ten year (i.e.  $k = 120$  months) real TIPS yields. Indexed TIPS offer a real future income stream that is known at the date of the bond's purchase.<sup>7</sup> Thus, observed yields reflect ex-ante interest rates in terms of fixed and known future purchasing power. With observed quotes  $E_t r_{t,k}$ , the investigation of (3.2) does not suffer from weak measurement of long run inflation expectations, systematic inflation expectation errors or omitted inflation risk premia. Under capital market integration, real TIPS yields equalize if agents *believe* that, until maturity, realized inflation differentials will materialize in exchange rate changes.

Allowing for transitory deviations from ex-ante relative PPP and UIP obtains

$$r_{t,k}^{(i)} - r_{t,k}^{(j)} = e_t, \quad (3.3)$$

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<sup>6</sup>In this chapter, we will use superscripts to indicate cross sectional members to avoid clumsy notation.

<sup>7</sup>Indexation of TIPS is subject to a small time lag (typically three months). According to the empirical evidence in Evans (1998), it is unlikely that this source of nominal exposure affects market prices.

where  $e_t \sim (0, \sigma_t^2)$  follows a stationary process with unconditional mean of zero and heteroskedastic innovations. Though equation (3.2) implies full convergence of RIRs many authors consider international credit markets as integrated under less restrictive conditions. For instance, Mishkin (1984) notes that one should not only address if a rejection of the RIP is statistically significant, but also if it is economically significant. Moreover, Goodwin et al. (1994) argue that, in presence of transaction costs or non-traded goods, RIRs differ across integrated capital markets. Lastly, a process of capital market integration might imply that conceptually non-stationary RIDs are characterized by systematic drift patterns that reflect a steady diminishing of differentials over time.<sup>8</sup>

### 3.2.2 Hypotheses

In this section we briefly state the hypotheses and sketch the implications for testing. As a starting point, consider a generalization of the model in (3.3),

$$r_{t,k}^{(i)} - r_{t,k}^{(j)} = c + d_1 t + d_2 t^2 + e_t, \quad e_t \sim (0, \sigma_t^2). \quad (3.4)$$

With reference to (3.4) the following hypotheses formalize distinct degrees of capital market integration.

1. The RIP holds in its strong form (H<sub>1</sub>):

Capital market integration implies that the law of one price holds internationally. Hence, expected real returns should be equal in equilibrium, i.e.  $c = d_1 = d_2 = 0$ .

2. The RIP holds in a weak form (H<sub>2</sub>):

In an integrated capital market, RIRs do not equalize in presence of transaction costs or non-traded goods. Moreover, the factual characteristics of the TIPS used to extract ex-ante RIRs are very similar across markets, but not identical. RIR convergence up to a (small) nonzero intercept, i.e.  $c \neq 0, d_1, d_2 = 0$ , indicates capital market integration which is in line with such stable frictions. In contrast, large persistent deviations from parity could indicate substantial currency risk premia that are at odds with market integration. As a decision rule, weak market integration is rejected if stationary RIDs exhibit risk premia

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<sup>8</sup>We thank an anonymous referee for this remark.



that appear economically meaningful. We reject  $H_2$  if a significant estimate  $\hat{c}$  exceeds 0.25 percentage points.<sup>9</sup>

3. Convergence via structural change ( $H_3$ ):

The continuous process of liberalization or a general decline in home biases might imply gradually shrinking RIDs. Respective adjustment patterns could be formalized in (3.4) by means of a deterministic trend that tends to vanish over time. With an initially positive (negative) RID, this would formally imply trend stationarity of RIDs and  $d_1 < 0$  and  $d_2 \geq 0$  ( $d_1 > 0$  and  $d_2 \leq 0$ ).

### 3.2.3 Methodology

To test hypotheses  $H_1$ ,  $H_2$  or  $H_3$  for a subsystem of the G7 we address the stationarity of selected RIDs. Since the convergence of CA-FR or CA-UK RIRs likely reflects their individual convergence to the US RIR, we confine the analysis to markets CA-US, UK-FR, FR-US and UK-US. Furthermore,  $H_1$  and  $H_2$  are investigated by means of bivariate cointegration tests. Building upon the results from pairwise market comparisons, finally, we test  $H_1$  in the four dimensional system.

Testing significance of risk premia and convergence via structural change ( $H_3$ ), notably,  $t$ -ratios of deterministic terms in ADF regressions follow distinct distributions if the considered process is integrated or not. Hence, respective significance tests rely on either Gaussian critical values (under stationarity) or their simulated counterparts (under the unit root hypothesis). However, as detailed in the empirical analysis, all main conclusions on capital market integration hold under common Gaussian critical values. To further enhance the reliability of inference, the empirical analysis takes into account that RIR innovations are heteroskedastic. In a couple of papers Cavaliere and coauthors (Cavaliere and Taylor 2007, 2008; Cavaliere, Rahbek and Taylor 2008) showed that neglecting non-stationary volatility can disturb the size features of widely applied unit root and cointegration diagnostics. Therefore, we derive  $p$ -values for respective test statistics by means of the wild bootstrap approach outlined in the aforementioned references. All main conclusions on (co)integration features are also supported by conventional inference.

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<sup>9</sup>Note that  $\hat{c} > 0.25$  is interpreted as economically meaningful since it is the smallest unit at which central banks change their target interest rates.

### Unit root tests

To infer on  $H_q$ ,  $q = 1, 2, 3$ , consider the ADF type regression (Dickey and Fuller 1979)

$$\Delta y_t^{(s)} = \gamma + \delta_1 t + \delta_2 t^2 + \phi y_{t-1}^{(s)} + \lambda_1 \Delta y_{t-1}^{(s)} + \dots + \lambda_p \Delta y_{t-p}^{(s)} + u_t^{(s)}, \quad u_t \stackrel{i.i.d.}{\sim} (0, \sigma_t^2). \quad (3.5)$$

For country pair  $s$ , the RID is  $y_t^{(s)} = r_{t,k}^{(i)} - r_{t,k}^{(j)}$ ,  $\{i, j\} \in \{\text{CA,US}; \text{UK,FR}; \text{FR,US}; \text{UK,US}\}$ . Testing the strong RIP ( $H_1$ ), deterministic terms are excluded from (3.5) to achieve maximum power. We diagnose  $H_1$  if the RID is stationary around zero ( $\phi < 0$ ). The weak RIP ( $H_2$ ) is diagnosed if the RID is stationary around a significant (but small) intercept estimate ( $\phi < 0, \gamma \neq 0$ ). On the condition that neither  $H_1$  nor  $H_2$  are empirically supported,  $H_3$  is diagnosed if scenarios ( $\phi < 0, \delta_1 < 0, \delta_2 \geq 0$ ) or ( $\phi < 0, \delta_1 > 0, \delta_2 \leq 0$ ) conform with the data rather than ( $\phi = 0$ ).

### Cointegration tests

In contrast to unit root tests, cointegration tests address a necessary condition for the international convergence of RIRs. In case of cointegration, one can subsequently test if the diagnosed long run relations are consistent with the RIP. Compared with unit root diagnostics cointegration methodology offers at least two benefits. Firstly, maximum likelihood based inference promises efficiency gains. Secondly, full dimensional modeling provides one further rank condition which is necessary (though not sufficient) for capital market integration. For simplicity of exposition consider a purely stochastic VECM representation

$$\Delta Y_t = \Pi Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \dots + \Gamma_p \Delta Y_{t-p} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \Sigma_t), \quad (3.6)$$

where  $Y_t$  denotes a  $N$ -dimensional vector of non-stationary RIRs and  $\Gamma_1, \dots, \Gamma_p$  are  $N \times N$  parameter matrices. Multivariate normality with time homogeneous covariance  $\Sigma$  is commonly assumed for maximum likelihood estimation. If the  $N \times N$  coefficient matrix  $\Pi$  has a reduced rank,  $\text{Rank}(\Pi) = g < N$ , there exist  $N \times g$  parameter matrices  $\alpha$  and  $\beta$  such that  $\Pi = \alpha\beta'$  and  $\beta'Y_{t-1}$  formalizes  $g$  stationary equilibrium relationships among country specific RIRs. Critical values for trace statistics (Johansen 1991) are determined by means of resampling (Cavaliere, Rahbek and Taylor 2008).

Bivariate cointegration analysis,  $N = 2$ , is performed for RIR tuples  $Y_t = Y_t^{(s)} = (r_{t,k}^{(i)}, r_{t,k}^{(j)})'$ , where  $\{i, j\} \in \{\text{CA,US}; \text{UK,FR}; \text{FR,US}; \text{UK,US}\}$ . In this context two capital markets are integrated in the strong form ( $H_1$ ) if RIRs share a common stochastic trend ( $g = 1$ ), and the cointegrating parameters conform with a long run parity

( $\beta' = (1, -1)$ ). The weak form of integration ( $H_2$ ) prevails if, in addition, a significant intercept of less than 0.25 percentage points enters the parity. Under full capital market integration the system of non-stationary RIRs,  $N = 4$ , is driven by one global stochastic trend. Accordingly, the cointegration rank is  $g = 3$ , and overidentifying restrictions formalize CA-US, FR-US and UK-US parities.

### 3.3 Convergence characteristics of ex-ante RIRs

The empirical analysis comprises three steps. After introducing the data, we conduct ADF tests for RIDs under alternative scenarios of RIR convergence ( $H_1$ ,  $H_2$  and  $H_3$ ). Then, pairwise cointegration tests for RIRs are discussed and parameter restrictions implied by  $H_1$  and  $H_2$  are tested. Finally, we turn to the full system and test for joint (i.e. 'global') capital market integration  $H_1$ . If not stated explicitly the discussion of diagnostic results refers to a nominal 5% significance level.

#### 3.3.1 Data and preliminary analysis

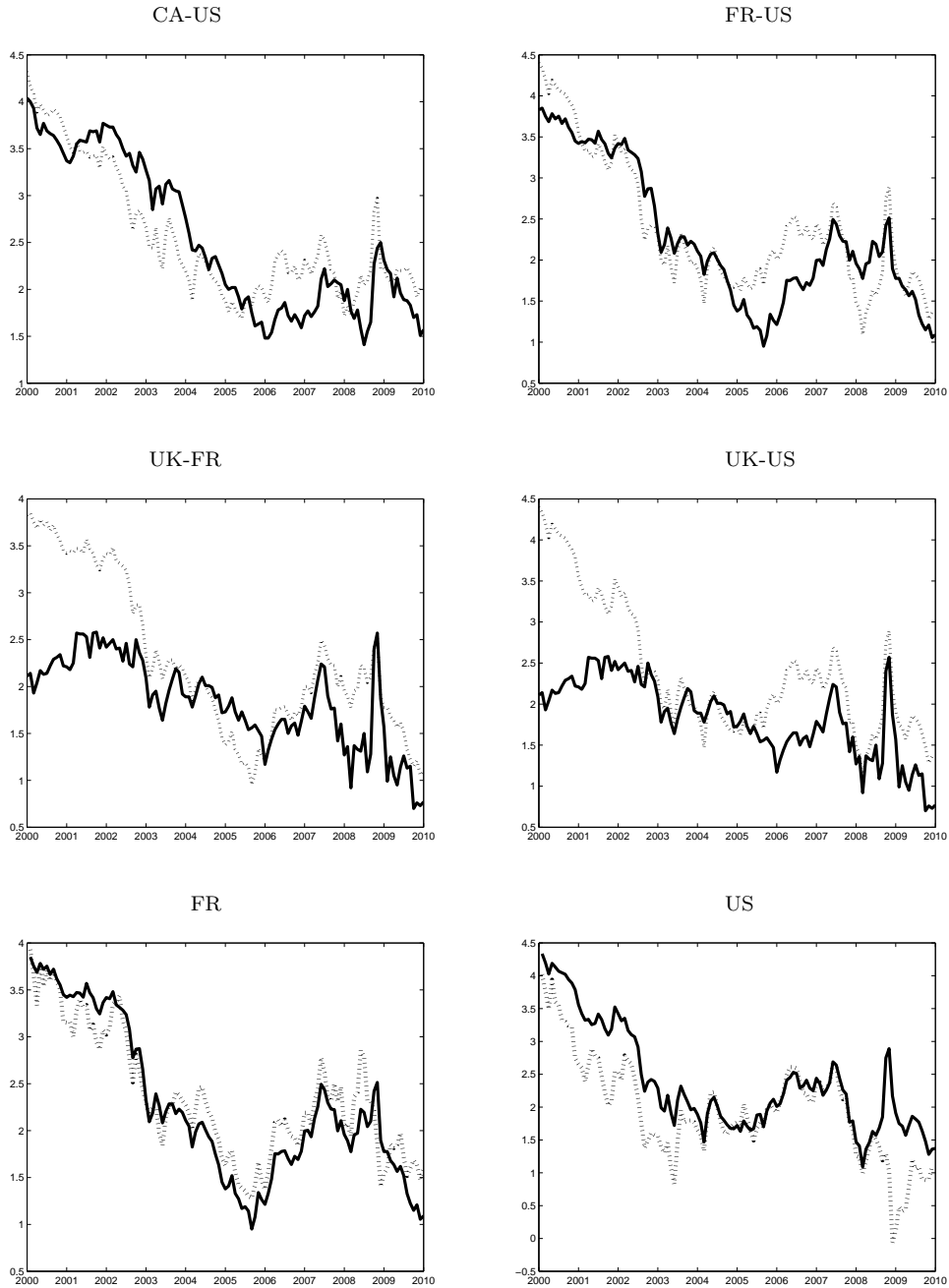
We analyze market quotes for CA, FR, UK and US real TIPS yields. The analysis is based on an update of the Herwartz and Roestel (2009) dataset used to investigate the international dependence of breakeven inflation rates. Taking liquidity of indexed long term government debt into account, ten year constant maturity RIRs are drawn whenever possible.<sup>10</sup> The Canadian RIR is approximated by means of the real government benchmark bond yield which refers to higher maturities. The Canadian government only issues real bonds that have a 30 year maturity. To examine CA-US capital market integration we contrast Canadian real benchmark bond yields to comparable US real benchmark bond yields constructed from 2029 and 2032 US TIPS. Further details on the data sources are given in Table 5.1 and associated RIRs are displayed in Figure 3.1. Visual inspection of the upper four panels of Figure 3.1 reveals a downward trend in RIRs that accords with the easing of monetary policy in response to the 2001 recession and the subsequent bond market 'Conundrum' (Greenspan 2004).<sup>11</sup>

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<sup>10</sup>From an economic perspective, 30 year interest rates should be used since, at this time horizon, relative PPP appears to be a reasonable assumption (Taylor and Taylor 2004). However, evidence in Section 3.4 suggests that, with respect to detecting long run convergence of RIRs, data quality is more crucial than maturity issues. Therefore we keep the adverse impacts of liquidity risks as small as possible rather than choosing assets with maturities that are fully consistent with empirical evidence on relative PPP.

<sup>11</sup>The 'Conundrum' has been associated with a general decline in term premia (Backus and Wright 2007).

Figure 3.1: Real interest rates



The upper four panels show long term ex-ante RIRs for CA, FR and the UK (solid) along with potential benchmark rates (dotted). The lower two panels show US and FR TIPS implied ex-ante RIRs with maturity  $k \approx 120$  months (solid) jointly with comparable conventional benchmark RIR approximations  $r_{t,120}^{(\bullet)} = \xi_{t,120}^{(\bullet)} - E_t[\pi_{t,120}^{(\bullet)}]$ , where  $\bullet \in \{\text{FR,US}\}$  (dotted).  $E_t[\pi_{t,120}^{(\bullet)}]$  is obtained from survey data.

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To further visualize linkages of local to world market RIRs, the upper four panels of Figure 3.1 show RIRs in a pairwise manner. Obviously CA, FR and UK RIRs tend to adjust to the RIRs of large partner economies, though spreads of either sign last up to three years.

Descriptive statistics of associated RIDs in Table 5.1 hint at a rough coincidence of CA-US and FR-US RIRs, since associated RIDs have unconditional means below 0.1% in absolute value while their standard deviation is 0.38%. Comparing the descriptive characteristics of *observable* long term RIDs and approximated long term differentials in Fujii and Chinn (2001), the former appear smaller with respect to their range and first order moments.

Table 3.1: Data

Type and source								
CA real 1-term	Datastream (CNBBRLT)							
FR OATI 2009, 2017, 2029	L'Agence France Trésor							
UK 10y real	Datastream (UKGR10Y)							
US TIPS 2009, 2011, 2012, 2030, 2032	Federal Reserve Bank of St.Louis							
US 10y real	Federal Reserve Bank of St.Louis							
Descriptive statistics								
RIRs				RIDs				
CA	FR	UK	US	CA-US	UK-FR	FR-US	UK-US	
2.544	2.298	1.834	2.392	Mean	-0.021	-0.455	-0.093	-0.558
0.806	0.819	0.467	0.805	S.D.	0.377	0.549	0.381	0.577
1.480	0.950	0.700	1.090	Min.	-0.630	-1.822	-0.862	-2.293
4.040	3.850	2.580	4.403	Max.	0.760	0.610	0.714	0.310

Notes: The upper panel documents TIPS and constant 10 year constant maturity RIRs used in this study. FR 10 year RIRs are approximated by linear interpolation of OATI 2009, 2017 and 2017, 2029 yields assuming a linear yield curve slope. Quotes on US constant maturity 10 year RIRs are available since 2003. Before, we approximate 10 year US interest rates by using 2009, 2011 and 2012 US TIPS yields. Moreover, the CA-US subsystem is investigated with the real yield of CA long term benchmark treasury bonds and US benchmark bond yields with a comparable maturity obtained from 2030 and 2032 TIPS. In the lower panel, the left (right) hand side documents descriptive statistics for RIRs (RIDs). S.D. is short for the empirical standard deviation.

Before 2003, the UK RIR is mostly lower than its French and US counterparts which might reflect temporary liquidity risk premia raised on TIPS markets other than UK. Findings in Shen (2006), Gürkaynak, Sack and Wright (2010) or Herwartz and

Roestel (2009) suggest liquidity risk premia to affect CA and US real yields before 2003. Notably, before 2003, the TIPS market volume in CA, FR or the US was lower than in the UK. Notwithstanding these technical features, the UK RIR seems to systematically adjust to the parity with the US rate.

Eventual liquidity risk premia<sup>12</sup> may complicate the analysis of RIR convergence. To illustrate the approximation accuracy of TIPS yields, the lower panels of Figure 3.1 display ten year FR and US real TIPS yields against corresponding RIR approximations obtained by subtracting survey based expected ten year inflation from ten year nominal interest.<sup>13</sup> Pointing to liquidity risk premia, US real TIPS yields exceed their conventional counterparts until 2003 while approximated and observed FR RIRs appear very similar throughout. After 2003, FR real TIPS yields tend to be slightly below their counterparts which implies that liquidity risk premia are close to zero under low inflation risk in FR. We conclude that liquidity risk premia are most likely transitory and, hence, should not severely bias inference under full convergence of RIRs. Moreover, as a consequence of non-synchronous risk patterns, potential premia are more likely to bias inference towards the hypothesis of market segregation.

Economic theory postulates that RIRs are determined by the marginal product of capital, and, hence, should be stationary around a positive constant. However, stochastic trending is formally confirmed by the ADF diagnostics in the left panel of Table 3.2. Since intercept estimates are insignificant throughout, we conclude that observed trending patterns do not reflect systematic drifts.

### 3.3.2 Stationarity of RIDs

Table 3.3 documents unit root test of RIDs. Particular test designs formalize *i*) convergence to zero, *ii*) convergence to a nonzero mean, and *iii*) gradual diminishing of RIDs. Explicitly testing for stationarity around zero, the strong RIP ( $H_1$ ) is confirmed for all RIDs, i.e. CA-US (with 10% significance), UK-FR, FR-US and UK-US. Generalizing the test regressions by means of intercept terms does not indicate significance of the respective parameters. Hence, the data are well in line with the parsimonious regression supporting market integration in the strong form. Lastly, trend stationarity of RIDs as formalized by  $H_3$  is not supported since trend parameter estimates,  $\hat{\delta}_1$  and  $\hat{\delta}_2$ , are insignificant.

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<sup>12</sup>We do not consider convergence properties of liquidity risk adjusted TIPS yields (Herwartz and Roestel 2009), since the respective adjustment factors depend on approximated long run inflation expectations.

<sup>13</sup>Arguably, this is one of the most precise ways to measure long run FR and US RIRs in conventional terms. Comparable data on long term ex-ante inflation is not available for CA and the UK.

Table 3.2: (Co)integration characteristics of RIRs

	Integration		Cointegration		
	ADF	$c$	$H_0$	eigenv.	trace
CA	-1.460 (.421)	0.035 (.682)	$g = 0$	0.195	52.18* (.009)
FR	-1.373 (.492)	0.025 (.793)	$g \leq 1$	0.114	26.61* (.056)
UK	-1.969 (.276)	0.137 (.343)	$g \leq 2$	0.078	12.34* (.083)
US	-2.240 (.141)	0.090 (.358)	$g \leq 3$	0.023	2.738 (.199)

ADF statistics and the corresponding intercept estimates are given in the left panel. The right panel documents trace test results for the four dimensional system of RIRs. Throughout, heteroscedasticity consistent  $p$ -values (Cavaliere and Taylor 2007, Cavaliere, Rahbek and Taylor 2008) are given in parentheses. In all ADF regressions, the selected lag length is 0 according to the SIC. The VECM is specified excluding deterministic terms in the cointegration space as suggested by the SIC. A lag of order 2 is chosen since highly significant first order autocorrelation prevailed for more parsimonious lag structures. For all test statistics, '\*' signals significance at the 5% level according to standard (i.e. non bootstrap) critical values.

An unreported cross check for stationarity of RIDs conducted for the 2004/1-2010/1 subperiod characterized by presumably sufficient market liquidity hints at  $H_1$  to prevail for UK-FR, CA-US and FR-US (the latter with 10% significance), while  $H_1$  to  $H_3$  are not supported for the case of the UK-US RID. Moreover, excluding the financial crisis from the sample (2000/1-2008/1) yields  $H_1$  to hold for UK-FR, FR-US (with 10% significance) and UK-US. Notably, rejecting non-stationarity within the shortened samples could be overly challenging under near integration. Empirical findings suggest that RIDs are stationary around zero if measured in terms of real TIPS yields. However, the adjustment is sluggish in general and unit root tests that include redundant deterministic terms fail to reproduce results on stationarity obtained from more restrictive designs.

### 3.3.3 Cointegration testing

We provide cointegration diagnostics for systems of RIRs testing  $H_1$  and  $H_2$ .<sup>14</sup> First, we check for cointegration among international RIRs. Then, we provide LR diag-

<sup>14</sup>Since univariate analysis did not indicate the existence of deterministic trends, trend terms are excluded from the cointegration space.

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nostics to test if the estimated cointegration relationships are consistent with  $H_1$  or  $H_2$ .

Table 3.3: Bivariate diagnostics

ADF tests for RIDs					Cointegration diagnostics for RIRs				
	CA-US	UK-FR	FR-US	UK-US		CA-US	UK-FR	FR-US	UK-US
ADF <sub>(1)</sub>	-1.860 (.069)	-2.369* (.005)	-2.294* (.043)	-2.919* (.002)	Tr <sub>(1)</sub>	13.41* (.039)	13.71* (.042)	13.94* (.033)	13.62* (.031)
ADF <sub>(2)</sub>	-1.860 (.267)	-2.477 (.109)	-2.306 (.227)	-3.111* (.021)	$\hat{\beta}$ (LR)	-0.918 (.127)	-0.799 (.365)	-0.891 (.100)	-0.847 (.297)
$\hat{\gamma}_{E-3}$	-2.659 (.959)	-18.16 (.642)	-5.302 (.903)	-28.34 (.519)		-	-	-	-
ADF <sub>(3)</sub>	-2.412 (.620)	-3.270 (.215)	-2.373 (.592)	-2.555 (.546)	Tr <sub>(2)</sub>	21.55* (.055)	18.52 (.082)	18.14 (.075)	18.54 (.087)
$\hat{\gamma}$	0.034 (.564)	-0.277* (.044)	0.028 (.623)	-0.097 (.551)	$\hat{\beta}$ (LR)	-1.344 (.244)	-0.399 (.029)	-1.166 (.446)	-2.119 (.516)
$\hat{\delta}_{1E-4}$	3.803 (.929)	-82.16* (.170)	-11.49 (.729)	26.41 (.682)	$\hat{c}$	1.305 (.028)	-0.827 (.002)	0.633 (.124)	2.581 (.077)
$\hat{\delta}_{2E-5}$	-0.288 (.944)	-6.029* (.242)	0.775 (.790)	-2.181 (.647)	$\tilde{c}$	0.218 (.074)	0.172 (.459)	0.272 (.026)	0.337 (.070)
Lag	0	0	1	0		0	0	2	0

The left hand side panel provides ADF statistics (heteroscedasticity consistent  $p$ -values, Cavaliere and Taylor, 2007) for RIDs. Three alternative test specifications are considered, namely without deterministic terms (ADF<sub>(1)</sub>), including a constant  $\hat{\gamma}$  (ADF<sub>(2)</sub>), and a constant and a time trend  $\hat{\delta}_1, \hat{\delta}_2$  (ADF<sub>(3)</sub>). The lag selection criterion is the SIC. '\*' signals significance at the 5% level according to standard (non bootstrap) critical values. The right hand side panel documents cointegration diagnostics (trace statistic, 'Tr', and LR test of the parity null hypothesis) for bivariate RIR vectors. The unrestricted cointegrating parameter is denoted  $\hat{\beta}$ , For trace statistics, heteroscedasticity consistent  $p$ -values (Cavaliere et al. 2008) are given in parentheses and '\*' signals significance at the 5% level according to standard critical values. Two alternative test specifications are considered, a cointegrating equation without deterministic terms (upper panel) and a specification with a constant in the long run equilibrium (lower panel). Intercept estimates  $\hat{c}$  ( $\tilde{c}$ ) are obtained from unrestricted and restricted ( $\beta = 1$ ) estimates, respectively. Lag selection is according to the SIC. If autocorrelation prevails, we choose the most parsimonious specification for which an LM diagnostic indicates absence of serial correlation.

First consider the set of bivariate VECMs ( $N = 2$ ). As documented in Table 3.3, results for bivariate cointegration tests with zero intercept are generally in line with



ADF diagnostics. Bivariate cointegration relationships are diagnosed for all considered capital market pairs, and restrictions implied by  $H_1$  are not rejected. Allowing for a non-zero intercept in the cointegration relation hardly changes the evidence for strong capital market integration. Bivariate cointegration among CA-US, UK-FR, FR-US and UK-US RIRs receives support at the 10% level. Intercept estimates are insignificant for the Northern American, Anglo American and the European subsystems and parameter restrictions cannot be rejected for CA-US, FR-US and UK-US RIRs. A small but significant intercept seems to prevail in the FR-US subsystem. As a potential reflection of sampling uncertainty, the parity among cointegrated UK-FR RIRs is rejected, which is at odds with the joint insignificance of respective restrictions applied to FR-US and UK-US RIRs. Thus, full convergence of real capital market rates,  $H_1$ , likely holds for the subsystems CA-US, UK-FR<sup>15</sup> and UK-US while the FR-US market might be better characterized by  $H_2$ . Interestingly, testing implications of ex-ante relative PPP, Bhatti and Moosa (1994) find strongest evidence against goods market integration for this particular market pair within a cross section of ten OECD economies.

In order to test if there is one 'global' capital market, we perform a full dimensional analysis ( $N = 4$ ). According to the SIC, we do not include an intercept in the cointegrating relations. Two lags are necessary to remove significant first order residual autocorrelation. Trace statistics given in the right hand side of Table 3.2 indicate the existence of  $g = 3$  cointegration relationships with 10% significance. Economically,  $g = 3$  holds if all RIRs are subject to the capital market of one major (global) economy, the US, say. Noting that evidence on bivariate cointegration also suggests  $g = 3$ , system features are discussed for the global trend model. For space considerations, Table 3.4 shows estimation results only for the overidentified VECM specification that incorporates the restrictions implied by  $H_1$ . Moreover, to improve the informational content of estimated error correction (EC) dynamics, insignificant EC coefficients have been sequentially set to zero. Presuming three cointegrating relationships and testing long run parities CA-US, FR-US and UK-US, the hypothesis of globalized markets accords with the data ( $p$ -value: 0.18). Most estimated EC coefficients have reasonable signs in the sense that deviations from the RIP are transitory. RIRs in CA, FR and the UK tend to close deviations from the US rate significantly. In general, the estimated adjustment speeds imply that eventual deviations have a half-life of approximately one year. Moreover, Canadian and UK RIRs respond to violations of

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<sup>15</sup>Since the UK-FR intercept term is estimated insignificant, we refer to the LR-statistic obtained in the parsimonious cointegration model.

the FR-US parity. To understand this result one might argue that if US assets lose global attraction, investors redirect their capital in favor of (similar) Anglo-American opportunities. US RIRs do not significantly adjust to deviations from parity relationships. Hence, the long term US RIR is diagnosed weakly exogeneous and therefore dominant in the long run. This result underpins US dominance and supports former evidence for estimated ex-ante RIRs (Al-Awad and Goodwin 1998).

Table 3.4: VECM estimates

Cointegration parameters and error correction dynamics									
	$r^{CA}$	$r^{FR}$	$r^{UK}$	$r^{US}$		$\Delta r^{CA}$	$\Delta r^{FR}$	$\Delta r^{UK}$	$\Delta r^{US}$
CI <sub>1</sub>	1	–	–	–1	EC <sub>1</sub>	–0.070 (.003)	–	–	–
CI <sub>2</sub>	–	1	–	–1	EC <sub>2</sub>	0.059 (.011)	–0.052 (.014)	0.045 (.044)	–
CI <sub>3</sub>	–	–	1	–1	EC <sub>3</sub>	–	–	–0.048 (.000)	–
Short term dynamics									
	$\Delta r^{CA}$	$\Delta r^{FR}$	$\Delta r^{UK}$	$\Delta r^{US}$		$\Delta r^{CA}$	$\Delta r^{FR}$	$\Delta r^{UK}$	$\Delta r^{US}$
CA(1)	–0.243 (.044)	–0.150 (.216)	–0.005 (.979)	–0.367 (0.042)	CA(2)	–0.039 (.742)	–0.268 (.026)	–0.255 (.200)	–0.256 (.154)
FR(1)	0.165 (.254)	–0.051 (.725)	–0.008 (.974)	–0.230 (.288)	FR(2)	–0.099 (.511)	0.068 (.654)	–0.004 (.988)	0.113 (.623)
UK(1)	0.030 (.709)	0.006 (.937)	–0.188 (.160)	0.048 (.691)	UK(2)	–0.002 (.984)	–0.061 (.441)	–0.261 (.047)	–0.136 (.251)
US(1)	0.141 (.226)	0.147 (.207)	0.222 (.254)	0.412 (.018)	US(2)	0.028 (.814)	–0.100 (.410)	0.259 (.187)	–0.039 (.823)

VECM parameter estimates for RIR systems and cointegration rank  $g = 3$ . The system is estimated excluding deterministic terms according to the SIC. Initially insignificant EC coefficients are restricted to zero,  $p$ -values are given in parentheses. Particular columns correspond to dependent variables, while rows indicate (lagged) explanatory variables and EC coefficient estimates. The LR statistic for testing overidentifying restrictions ( $p$ -values) implied by long run CA-US, FR-US and UK-US parities is 4.839 (.184). System residual diagnostics ( $p$ -values) are: White heteroscedasticity test (.000), Jarque-Bera test for normality (.000), LM test against first order serial correlation (.611).

### 3.4 A note on evidence in related studies

One may argue that previous investigations of RIR convergence are subject to a trade-off between data quality and maturity: On the one hand, the central building blocks of the RIP, namely the UIP and the ex-ante relative PPP, hardly hold for short term horizons. As RIRs are typically obtained by subtracting realized inflation from short term interest rates (e.g. Goodwin and Grennes 1994, Chinn and Frankel 1995, Ferrera and Ledesma 2007), this caveat applies to most studies. On the other hand, RIR approximation accuracy strongly decreases at higher maturities due to inflation risk premia and imprecise measures of expected long run inflation. Since quotes on survey-based long run inflation expectations are rare, inflation compensation of long run nominal interest rates has been approximated by means of ten year realized forward inflation (Fujii and Chinn 2001), one year realized forward inflation (Obstfeld and Shambaugh 2002) or model based inflation forecasts (Fujii and Chinn 2001, Al-Awad and Goodwin 1998). Arguably, all these statistics are more likely subject to systematic biases than corresponding short run measures.

While evidence at hand suggests that the RIP holds if RIRs are measured in terms of real TIPS yields, related studies tend to reject the RIP. To investigate if this discrepancy reflects maturity or measurement issues we conduct ADF tests for RIDs calculated from distinct short and long run RIR measures. Evidence for non-stationarity of approximated short term (long term) RIDs is seen to support the suggestion that the maturity (measurement) issue is of particular importance. Lastly, overall convergence for all alternative RID measures is seen to indicate that market integration has increased in the most recent decade. We approximate RIRs in three alternative ways.

1. Firstly, short term RIRs are obtained by subtracting the three month forward consumer price inflation (CPI) rate from the three month nominal interest rate, i.e.

$$r_{t,3}^{(i)} = \xi_{t,3}^{(i)} - \pi_{t,3}^{(i)}, \quad (3.7)$$

where  $\pi_{t,3}^{(i)} = \ln \text{CPI}_{t+3}^{(i)} - \ln \text{CPI}_{t-9}^{(i)}$ . Given moderate inflation and price stickiness, inflation expectation errors and potential premia for inflation risk should be small for an investment horizon of three months, say.

2. Secondly, long term RIRs are obtained by subtracting survey based measures of expected inflation from long run nominal interest rates. In particular, CA-US, UK-FR and UK-US RIDs are derived from RIRs

$$r_{t,120}^{(i)} \approx \xi_{t,120}^{(i)} - E_t \pi_{t,12}^{(i)}, \quad (3.8)$$

where  $\xi_{t,120}^{(i)}$  is the ten year nominal interest rate and  $E_t\pi_{t,12}^{(i)}$  is the inflation rate expected to prevail during the next year. Since survey based expected long run inflation rates are available for the European Monetary Union (i.e. FR) and the US, we investigate FR-US market integration by means of long term RIRs

$$r_{t,120}^{(i)} \approx \xi_{t,120}^{(i)} - E_t\pi_{t,120}^{(i)}, \quad (3.9)$$

where  $E_t\pi_{t,120}^{(i)}$  are survey based ten year inflation expectations.<sup>16</sup> Notably, quotes on expected inflation for CA, FR and the UK are only available at the quarterly frequency. Thus, empirical results on convergence of approximated RIRs building on survey based ex-ante inflation are obtained from quarterly data for the period 2000q1 to 2009q4 which is suppressed in the notation to avoid confusion.

3. Lastly, long term RIRs are obtained in the spirit of Fujii and Chinn (2001). However, subtracting ex-post ten year future inflation from ten year nominal interest is not feasible for our sample period 2000/1 – 2010/1. Therefore, the long run steady state of inflation is determined by means of HP filtering (Hodrick and Prescott 1997). Formally, respective RIRs read as

$$r_{t,120}^{(i)} \approx \xi_{t,120}^{(i)} - \tilde{\pi}_t^{(i)}, \quad (3.10)$$

where  $\xi_{t,120}^{(i)}$  is the ten year nominal interest rate and  $\tilde{\pi}_t^{(i)}$  is the HP implied steady state of realized CPI inflation  $\pi_t^{(i)}$  in month  $t$ .

ADF test results for the three alternative RID measures are shown in Table 4.4. In line with the mixed evidence on market integration in related studies, our findings appear inconclusive when RIRs are measured conventionally. Short term RIDs are diagnosed stationary around zero for CA-US and UK-FR, while FR-US and UK-US RIDs are found to be non-stationary. In turn, ADF tests for differentials between approximated high maturity RIRs according to (3.8) and (3.9) indicate UK-US stationarity around an economically significant risk premium while full convergence might only hold for the FR-US RID. Furthermore, defining RIRs according to (3.10), we diagnose  $H_1$  and  $H_2$  for the UK-FR and CA-US markets with 10% significance, respectively. Hence, under weak approximation accuracy of RIRs or using short term

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<sup>16</sup>US survey based expectations refer to ten years ahead, while EMU expectations refer to five years only. Noting that EMU survey based expectations hardly show any variation, one may assume that five and ten year EMU expected inflation rates coincide.

variables, empirical tests might fail to confirm a de-facto state of capital market integration. In particular the weak approximation of RIRs seems to pose a more severe problem for inference since full convergence holds for some short term RIDs, and for the survey based long term FR-US RID given in (3.9).

### 3.5 Concluding remarks

We investigate capital market integration as implied by the convergence of *observed* ex-ante long term TIPS yields for the period 2000/1-2010/1 and markets CA, FR, the UK and the US. Unit root analysis suggests that all RIDs are stationary around zero. Evidence from bivariate cointegration testing points at the prevalence of cointegration for all markets. Moreover, observed cointegration patterns cannot be rejected to reflect pairwise integration of CA-US, UK-FR, FR-US and UK-US capital markets. Full dimensional cointegration analysis suggests that all RIRs are individually linked to the US counterpart. Tests of overidentifying parameter restrictions formally confirm the integration of CA, FR, UK and US capital markets. Moreover, though a growing financial influence of the European economies has been highlighted in Ehrmann and Fratzscher (2004), estimated adjustment dynamics suggest that the (global) cost of capital is determined in the US. Addressing the mixed evidence on the RIP in related studies, we show that, for the same time period, real money market rates or (weakly) approximated long term RIRs hardly converge. Hence, the failure of short run ex-ante relative PPP and short run UIP, and presumably weak approximations of long term RIRs might have favored empirical rejections of full capital mobility in former studies.

## Chapter 4

# Monetary Independence I: Interest Rate Autonomy

### 4.1 Motivation

The macroeconomic trilemma (Obstfeld, Shambaugh and Taylor, 2005) suggests that the openness of capital markets and the flexibility of FX rates jointly determine the extent to which small economies are able to isolate domestic interest rates from world interest rates. For small open economies, achieving a certain degree of autonomy over domestic interest rates is only possible by accepting some loss of control over domestic exchange rates and vice versa. Monetary dependence matters for small open economies that stabilize their FX rates to currencies of larger partner countries by means of FX market interventions. It further matters for economies that are subject to the notion of 'Fear of Floating' (Calvo and Reinhart, 2002). These formally 'floating' economies tend to follow world market interest rate movements to prevent strong adjustments of exchange rates. In both cases, however, the autonomy over domestic interest rates is determined by the extent of FX flexibility that monetary authorities are willing to tolerate. Hence, every small open economy that allocates nonzero weight to some FX rate target has to accept some loss of control over domestic interest rates. The possibility to choose among a number of distinct domestic policy strategies is the lead argument to allow for a floating currency (Calvo and Mishkin, 2003). In this respect, a comprehensive empirical assessment of this trade-off is important for a sensible evaluation if (and to what extent) a given economy should stabilize FX rates. Recent empirical evidence casts doubt if actual gains in monetary independence are substantial. Frankel, Schmukler and Servén (2004) suggest that, similar to pegs, full transmission of global interest rates also holds for free floating regimes, at least in the long run. According to Frankel et al. (2004), only three economies were able

to set their own interest rates over the 1990s, namely Germany, Japan and the US. Presuming that the decision to either pursue a peg or a nonpeg (i.e. to allow for some variation in FX rates) factually matters for the monetary independence of small economies, Shambaugh (2004) and Obstfeld et al. (2005) find interest rates of pegs to follow the base countries' rates closer in comparison with nonpegs.

The issue of monetary independence has been investigated by means of rather restrictive econometric models. The conventional approach is conditional on a highly stylized classification with respect to an observed or declared status of FX rate flexibility. Then, assuming parameter homogeneity within class specific subsamples, (unbalanced) pooled panel regression models for interest rate transmission are estimated. Available empirical evidence might suffer from the following shortcomings. Firstly, a conventional de facto classification of exchange rate flexibility such as pegs vs. nonpegs (Shambaugh 2004, Obstfeld et al. 2005) could be too restrictive. Moreover, the assumption of parameter homogeneity within rather general groups is likely violated and might lead to biased panel estimates. In fact, many economies do neither pursue pure pegs nor pure floats (Fischer 2001, Yeyati and Sturzenegger 2005) and these policies may also continuously change over time. To assess the macroeconomic trilemma, one should fully exploit the apparent time and country specific heterogeneity that prevails in the data, rather than relying on restrictive pooled panel regressions. Moreover, a lack of selectivity among country specific policy options with regard to FX flexibility limits the scope of empirical results for monetary policy advice. Secondly, former studies primarily focus on the impact of exchange rate stability on monetary independence, while the joint impact of capital mobility and exchange rate flexibility has not been sufficiently highlighted yet. While Frankel et al. (2004) indirectly consider capital mobility by modeling time specific subsamples, Shambaugh (2004) provides estimates for interest rate transmission conditional on capital mobility (classified according to the existence and absence of capital barriers) and currency regime type (peg vs. nonpeg). However, as a consequence of country specific currency risk premia, capital mobility might substantially differ across economies with liberalized capital markets, especially in case of nonpegs. Distinguishing between absence of capital barriers and capital mobility might not be essential for a comparison of transmission dynamics among (credible) pegs and nonpegs, but it matters for a comprehensive, continuous assessment of the macroeconomic trilemma. For instance, consider the case of economies following a reference countries' interest rates to prevent fluctuations of its exchange rate. For a given degree of FX rate variability, country specific independence crucially depends on the extent of capital mobility, as it determines the

strength of reactions of FX rates to unilateral interest rate changes. Hence, cross correlation of interest rate changes should continuously increase with capital mobility for all economies that attribute nonzero weight to FX targets. In a general model, therefore, interest rate transmission should be evaluated conditional on a continuum of representative states characterized by country specific factual FX flexibility and capital mobility. Thirdly, comovements among international interest rates have been assumed to reflect a loss of monetary autonomy in general, although the prevalence of real business cycle linkages suggests that central banks could independently choose a similar stance of monetary policy. Therefore, one should control for deviations from potential targets of inflation or output in an empirical model. Lastly, an analysis of monetary independence should be conducted within an uniform empirical model. Respecting the latter premise is not trivial since international interest rates might be characterized by heterogeneous (co)integration characteristics. To circumvent this difficulty, Frankel et al. (2004), Shambaugh (2004) and Obstfeld et al. (2005) employ an econometric approach suggested by Pesaran, Shin and Smith (2001). This methodology allows for uniform model estimation irrespectively if the underlying variables are  $I(0)$  or  $I(1)$ . However, it somehow exchanges the problem of model choice against the selection of critical values for inferential purposes which depend on assumptions concerning the (uniform) integration order that characterizes the data.

To address these issues, the empirical model in this work is implemented in the framework of flexible semiparametric functional coefficient models (Cai, Fan and Yao 2000, Herwartz and Xu, 2009). It allows interest rate transmission parameters to be estimated as a (nonlinear) function of both the current state of measurable FX variability and capital mobility. Moreover, country specific domestic fundamentals such as gaps of output, inflation and FX rates are included to identify if dependence of interest rates is induced by international capital flows or reflects independent policy steps under real economic linkages. Using a slightly modified concept of long run monetary dependence, we allow the domestic interest rates to adjust to a flexible (i.e. local) steady state obtained by Hodrick-Prescott (HP) filtering the interest rate differential.<sup>1</sup> Since deviations from HP implied steady states are stationary by construction, standard inferential tools apply irrespectively if cointegration features a system of international interest rates or not. The considered panel comprises quarterly data for 20 developed economies collected after the great moderation period 1987Q1-2008Q2.

Assessing the full trade-off among FX stability, capital mobility and monetary au-

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<sup>1</sup>With respect to the uncovered interest rate parity, the HP trend should reflect general differences in inflation rates (implying a permanently expected change in the domestic FX rate) or persistent premia for currency risk



tonomy in a general model improves upon earlier studies since it allows to quantify the presumed effects of any given reduction in country specific exchange rate flexibility on monetary dependence. Since monetary authorities might not realize substantial gains in independence when tolerating high variability in FX rates, this aspect is of natural interest for policy advice. Moreover, monetary authorities might be interested in the scope that is left for influencing domestic interest rates given the average extent of FX volatility they are willing to accept.

To preview the empirical results of this work, we confirm the implications of the trilemma to hold throughout. Monetary autonomy appears completely lost for economies that feature low FX rate flexibility and high capital mobility, while economies with rather immobile capital and flexible FX rates are least affected by interest rate transmission. For the latter, however, transmission still appears to be quite high. Formally testing the implications of the macroeconomic trilemma, we strongly reject invariance of interest rate transmission with respect to capital mobility and FX rate flexibility for all provided semiparametric model specifications. Notably, our results appear robust to alternative measures of FX rate flexibility, capital market openness and long run steady states among interest rates. Moreover, we provide evidence that ignoring the impact of domestic fundamentals leads to upward biased estimates of interest rate transmission especially for more flexible regimes. Thus, former evidence suggesting that floating implies only moderate advantages in monetary independence might arise from omitted variable biases.

The remainder of this chapter is organized as follows: In the subsequent section, we introduce our empirical model, which, in a first step, is estimated by means of conventional parametric panel models allowing for fixed effects and country specific interest rate rules. In section 4.3 we explain how the model is implemented in the functional coefficient framework, motivate the employed state variables and provide semiparametric estimation results. Furthermore, we discuss the findings and implications for monetary policy. Section 4.4 concludes. Technical details about estimation and inference in semiparametric models, as well as several robustness checks are given in the appendix.

## **4.2 Monetary independence - benchmark approaches**

As a starting point, we formalize interest rate transmission in a conventional (parametric) framework to obtain a (descriptive) assessment of data inherent features that allows a meaningful comparison with conclusions available from the literature. Firstly,

the empirical model is introduced, followed by a brief description of the regime classification procedures applied to measures of FX variability and capital mobility. After providing some information on the data set, parametric fixed effects panel regressions for each regime are conducted and empirical (benchmark) results are discussed.

### 4.2.1 Single country regressions

The implications of FX stability on monetary dependence have been investigated by comparing instantaneous transmission of interest rates across different FX regimes. Frankel et al. (2004) use monthly interest rate levels and unbalanced panel models with fixed effects, while Shambaugh (2004) considers first differences in yearly interest rates and pooled model estimation instead. We generalize the panel model in Shambaugh (2004) by additionally considering domestic fundamentals and a 'pseudo error correction term' that allows for long term adjustments to a time varying steady state. In line with Shambaugh (2004), the analysis builds on lower frequency data to minimize problems associated with heterogeneous time lags in short term adjustment dynamics. For given presample values consider single country regression models of the following form:

$$\begin{aligned}\Delta\xi_{it} &= \alpha_{i1}\Delta\xi_{jt} + \alpha_{i2}(\xi_{it-1} - \xi_{jt-1} - \phi_{ijt-1}^{hp}) + \beta_{i0} + \beta_{i1}\tilde{\pi}_{it-1} + \beta_{i2}\tilde{y}_{it-1} + \beta_{i3}\tilde{s}_{ijt-1} + e_{it}, \\ y_{it} &= \mathbf{x}'_{it}\alpha_i + \mathbf{z}'_{it}\beta_i + e_{it}, \quad t = 1, \dots, T, \quad i = 1, \dots, N,\end{aligned}\tag{4.1}$$

where the time and cross section dimension of the sample are denoted as  $T$  and  $N$ , respectively,  $\mathbf{x}'_{it} = [\Delta\xi_{jt}, (\xi_{it-1} - \xi_{jt-1} - \phi_{ijt-1}^{hp})]$  and  $\mathbf{z}'_{it} = [1, \tilde{\pi}_{it-1}, \tilde{y}_{it-1}, \tilde{s}_{ijt-1}]$ . Specifically,  $\xi_{it}$  is the (quarterly) short term domestic market interest rate in country  $i$ ,  $\phi_{ijt-1}^{hp}$  is the steady state interest rate differential between country  $i$  and reference country  $j$  obtained by HP filtering of the nominal interest rate differential  $\xi_{it-1} - \xi_{jt-1}$ . For benchmarking purposes, we use either German or US interest rates, where the German rate is the reference for all European countries (except Germany and the UK), and the US rate is considered for all other economies (including Germany and the UK). This choice reflects the commonly held view that most European interest rates are predominantly subject to German interest rates, while the US monetary policy tends to dominate world interest rates (Katsimbris and Miller, 1993, Kirchgässner and Wolters, 1993, Hassapis, Pittis and Prodromidis 1999). Turning to domestic fundamentals, the deviations from the steady states of domestic inflation, output and exchange rates are given by  $\tilde{\pi}_{it-1}$ ,  $\tilde{y}_{it-1}$  and  $\tilde{s}_{ijt-1}$ . Throughout, HP implied

steady states are recursively evaluated (and therefore 'observable').<sup>2</sup> As suggested by Hodrick and Prescott (1997) we use a smoothing parameter of  $\lambda = 1600$  for quarterly data. For exact definitions of the variables see table 5.1.

Obviously, the specification in (4.1) is similar to an error correction model restricted to a long run interest rate parity that holds up to an additional time varying intercept  $\phi_{ijt-1}^{hp}$  which is determined outside the model. In the framework of the uncovered interest rate parity,  $\phi_{ijt-1}^{hp}$  should reflect steady states of i) expected FX rate changes as e.g. implied by differences in (expected) inflation rates or ii) persistent risk premia. We choose a flexible local steady state since risk premia and cross country spreads in inflation expectations may trend in a persistent manner. Hence, imposing conventional deterministic terms such as an intercept or a time trend to enter a cointegrating relation might not be appropriate. Moreover, regression (4.1) is not subject to inferential issues associated with nonstationary (and eventually not cointegrated) variables, since deviations from the Hodrick-Prescott (HP) filtered interest rate differential are stationary by construction. As a robustness check we also provide estimates of conventional error correction dynamics in the appendix. Lastly, controlling for domestic fundamentals such as the deviations from HP-implied steady states of inflation, output and exchange rates in model (4.1) helps to rule out spurious evidence on monetary dependence that could be induced by real economic linkages. HP filtering is necessary since inflation rates and FX rates tend to be non-stationary. When considering domestic fundamentals  $\tilde{\pi}_{it-1}$ ,  $\tilde{y}_{it-1}$  and  $\tilde{s}_{ijt-1}$ , notably, we do not attempt to estimate the 'true' monetary policy reaction function for each country in the panel since this appears hardly feasible. Our aim is rather to improve the estimation of interest rate transmission compared with the case where effects of business cycle synchronization remain unconsidered at all. According to the model in (4.1), monetary independence implies that an economy is able to maintain isolated interest rate adjustments from a given equilibrium interest rate differential (as it is implied by HP trends) to influence output or inflation. Thus, the impact of international transmission (parameterized by  $\alpha_{i1}$  and  $\alpha_{i2}$ ) on the domestic interest rate should be insignificant or at least small in absolute terms. Moreover, the response of local interest rates to domestic inflation and output (parameterized by  $\beta_{i1}$  and  $\beta_{i2}$ ) should be (significantly) positive whenever central banks adopt monetary strategies in the spirit of the Taylor rule. A positive parameter estimate  $\hat{\beta}_{i3}$ <sup>3</sup> might reflect that monetary authorities adjust domestic interest rates in order to stabilize FX rates, and thus, their policy is not independent in

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<sup>2</sup>To guarantee sensible HP gaps at the beginning of the sample (1987Q1) we use presample information, i.e. a gap in  $t$  is derived by filtering data for the subperiod  $t^*, \dots, t$ , where  $t^* \ll 1987Q1$ .

<sup>3</sup>Note that FX rates are defined in direct quotation.

a stricter sense. In turn, loss of autonomy occurs if  $\alpha_{i1}$  (or  $\alpha_{i2}$ ) are estimated significantly positive (negative) and large in absolute value. Moreover, under high exposure to foreign interest rates, estimates of  $\beta_{i1}, \beta_{i2}$  and  $\beta_{i3}$  might be insignificant since, in such a setting, monetary authorities could fail to adjust interest rates with respect to domestic goals. For our purposes, however, the main emphasis lies on transmission parameters  $\alpha_{i1}$  and  $\alpha_{i2}$  since insignificance of  $\beta_{i1}, \beta_{i2}$  and  $\beta_{i3}$  might also signal that countries do not (try to) set domestic interest rates in response to deviations from output, inflation or FX rate targets. Nevertheless, a brief discussion of associated parameter estimates is provided for completeness. Partialling out  $\mathbf{z}_{it}$  from (4.1) yields

$$\check{y}_{it} = \check{\mathbf{x}}_{it}'\alpha_i + \check{e}_{it}, \quad (4.2)$$

where  $\check{y}_i = M_i y_i$ ,  $\check{\mathbf{x}}_i = M_i \mathbf{x}_i$ ,  $\check{e}_i = M_i e_i$  and  $M_i = (I_i - \mathbf{z}_i(\mathbf{z}_i' \mathbf{z}_i)^{-1} \mathbf{z}_i')$ , with  $\mathbf{z}_i = [\mathbf{z}_{i1}, \dots, \mathbf{z}_{iT}]'$ . For panel estimation of equation (4.2), we allow that interest rate transmission parameters  $\alpha_1$  and  $\alpha_2$  vary according to observable heterogeneity measured in terms of FX rate flexibility and capital market integration over countries  $i$  and for time  $t$ . Different assumptions concerning the pattern of heterogeneity give rise to both conventional panel regressions with fixed effects and country specific fundamentals (sections 4.2.2 and 4.2.4) as well as more flexible semiparametric functional coefficient panel regressions (section 4.3).

## 4.2.2 State definitions and panel estimation

Let observable country specific FX variability  $\tilde{\psi}_{it}^{fx}$  and capital (im)mobility  $\tilde{\psi}_{it}^{cap}$  in quarter  $t$  be approximated as

$$\tilde{\psi}_{it}^{fx} = \sum_{m \in t-1} (\Delta \ln s_{ij,m})^2 \quad \text{and} \quad \tilde{\psi}_{it}^{cap} = (r_{it-1} - r_{jt-1})^2, \quad (4.3)$$

respectively, where  $s_{ij,m}$  is the monthly price for one unit of reference currency  $j$  in terms of domestic currency  $i$  and  $r_{it} - r_{jt}$  is the quarterly real interest rate differential. Specifically,  $r_{it} = \xi_{it} - \pi_{it}^e$  is defined as the three month nominal interest rate  $\xi_{it}$  deflated by the three month expected rate of inflation  $\pi_{it}^e$  in time  $t$ . To avoid endogeneity issues we throughout approximate  $\pi_{it}^e$  by annualized realized consumer price inflation,  $\pi_{it}^e \approx \ln(CPI_{it}) - \ln(CPI_{it-4})$ , implicitly assuming static instead of rational expectations. The FX rate regime in country  $i$  at time  $t$  is classified as relatively 'flexible' ('inflexible') if the current FX volatility is below (above) the corresponding global i.e.

cross sectional median<sup>4</sup>,  $Med(\tilde{\psi}_t^{fx})$ . Moreover, to capture potential strengthening of capital market integration over time, it is distinguished if the capital mobility in time  $t$  is below or above the country specific median reference level<sup>5</sup>,  $Med(\tilde{\psi}_i^{cap})$ . Formally, we have

$$\tilde{\psi}_{it}^{fx} = \begin{cases} \tilde{\psi}_{it}^{fx,h} & \text{if } \tilde{\psi}_{it}^{fx} \geq Med(\tilde{\psi}_t^{fx}) \\ \tilde{\psi}_{it}^{fx,l} & \text{otherwise} \end{cases} \quad \text{and} \quad \tilde{\psi}_{it}^{cap} = \begin{cases} \tilde{\psi}_{it}^{cap,h} & \text{if } \tilde{\psi}_{it}^{cap} < Med(\tilde{\psi}_i^{cap}) \\ \tilde{\psi}_{it}^{cap,l} & \text{otherwise.} \end{cases}$$

Following the panel approach of Shambaugh (2004), we conduct regime specific pooled estimates of model (4.2) giving rise to the specification

$$\check{y}_{it} = \check{\mathbf{x}}_{it}'\alpha + \epsilon_{it}, \quad \forall \{i, t\} \in (\tilde{\psi}^{cap,\bullet}, \tilde{\psi}^{fx,\circ}), \quad (4.4)$$

where  $\bullet, \circ \in \{l, h\}$ . Panel model (4.4) implicitly includes country specific fixed effects accounting e.g. for long term moderation processes. To benchmark our model with alternative specifications that have been used earlier in this context, we complementary apply a classification scheme similar to that in Shambaugh (2004), where a country  $i$  in quarter  $t$  is labeled 'peg' if its (monthly) exchange rate has been within 2% bands over the last 12 months. To measure capital mobility, Shambaugh (2004) considered (non)existence of capital barriers. Since capital markets had been liberalized over the vast majority of 'small' developed economies until 1987, we do not take capital controls explicitly into account.

### 4.2.3 Data and variable definition

This study builds on quarterly cross sectional data for the time period 1987q1-2008q3 for 20 'small' economies, namely Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Ireland, Japan, South Korea, Netherlands, New Zealand, Norway, Singapore, Spain, Sweden, Switzerland and the UK. The choice is motivated by data availability. Outlying observations quoted in the context of international crises (e.g. European currency crisis 1992/1993, the Asian financial crisis 1997) are excluded from the sample as well as quotes related to excessive interest rate volatility. In particular, strong fluctuations in South Korean interest rates observed before 1999 completely dominated sample information and therefore rendered panel

<sup>4</sup>It has been argued (Yeyati and Sturzenegger 2005) that FX volatility might not only reflect FX flexibility, but also common exposure to shocks. It is noteworthy that defining FX volatility with respect to a global volatility benchmark might reduce this effect to some extent.

<sup>5</sup>Note that an evaluation with respect to the cross sectional median of capital mobility  $Med(\tilde{\psi}_t^{cap})$  would remove global time trends as e.g. implied by the process of financial globalization.

coefficient estimates hardly representative. Interest rates are three month money market rates, or, if not available, three month treasury bill rates. Spot FX rates are throughout determined in direct quotation. To measure inflation we use annualized CPI inflation. Annualized real growth of output is approximated by means of GDP data whenever possible. If corresponding time series are not available, annualized real growth of industrial production is used instead. Table 4.1 provides more detailed information about data sources, variable definitions and removed observations.

Table 4.1: Variable definitions, data sources and removed observations

Variable	Definition	Source	Removed Observations
$\xi_{it}$	3 Month MM/TB Rate	IFS	Bel 1992Q2-1993Q4*
$\pi_{it}, \pi_{it}^e$	$\ln(CPI_{it}) - \ln(CPI_{it-4})$	IFS	Den 1992Q2-1993Q4*
$y_{it}$	$\ln(GDP_{it}) - \ln(GDP_{it-4}) - \pi_{it}$	Datastream	FR 1992Q2-1993Q4*
	$\ln(IP_{it}) - \ln(IP_{it-4}) - \pi_{it}$	Datastream	IT 1992Q2-1993Q4*
$s_{ij}$	Price currency $i$ / Price currency $j$	Datastream	Ko 1987Q1-1999Q1+
$r_{it}$	$\xi_{it} - \pi_{it}$	-	Nor 1992Q2-1993Q4*
$\tilde{\pi}_{it}$	$\pi_{it} - HP(\pi_{it}, \lambda = 1600 I_t)$	-	NZ 1987Q1-1988Q3+
$\tilde{y}_{it}$	$y_{it} - HP(y_{it}, \lambda = 1600 I_t)$	-	Swe 1992Q2-1993Q4*
$\tilde{s}_{ij,t}$	$s_{ij,t} - HP(s_{ij,t}, \lambda = 1600 I_t)$	-	Fin 1992Q2-1993Q4*
$\phi_{ij,t}^{hp}$	$\xi_{it} - \xi_{jt} - HP(\xi_{it} - \xi_{jt}, \lambda = 1600 I_t)$	-	Ir 1992Q2-1993Q4*

Variable definitions and sources are given in the left panel of table 4.1. In particular, the notion  $HP(y_{it}, \lambda = 1600|I_t)$  refers to the HP implied trend of  $y_{it}$  using a smoothing parameter of  $\lambda = 1600$  and only quotes on  $y_{it}$  available until  $t$ . The right hand side panel lists observations discarded from the estimation procedure. Time periods marked with '\*' denote exclusion due to unusual fluctuations in corresponding interest rates in the line of the European currency crisis. Periods marked with '+' denote exclusion due to excess interest rate volatility.

#### 4.2.4 Results

Table 4.2 reports panel estimates for the model in equation (4.2) according to distinct regimes of FX flexibility and capital mobility.

As implied by a parameter estimate of  $\hat{\alpha}_1 = 0.81$ , changes in world interest rates  $\Delta\xi_{jt}$  are almost instantaneously reflected in interest rate changes of corresponding partner economies given that exchange rate volatility is low ( $\tilde{\psi}^{fx,l}$ ) and capital mobility is high ( $\tilde{\psi}^{cap,h}$ ). Integrated economies with low FX rate variation are therefore hardly able to adapt domestic interest rates in response to their individual positions over

the business cycle. Short run transmission to economies with low FX flexibility but less mobile capital ( $\tilde{\psi}^{fx,l}, \tilde{\psi}^{cap,l}$ ) is weaker ( $\hat{\alpha}_1 = 0.65$ ). At a nominal level of 5%, we do not diagnose significance of error correction dynamics  $\hat{\alpha}_2$  for these two regimes. Evidence in the subsequent chapter suggests considerable parameter heterogeneity to prevail within both groups. Consequently, the parameter heterogeneity not taken into account by this panel regressions appears to increase estimation uncertainty and results in imprecise panel estimates.

Table 4.2: Panel estimates for model 4.2

	All	$\tilde{\psi}^{fx,l}, \tilde{\psi}^{cap,h}$	$\tilde{\psi}^{fx,l}, \tilde{\psi}^{cap,l}$	$\tilde{\psi}^{fx,h}, \tilde{\psi}^{cap,h}$	$\tilde{\psi}^{fx,h}, \tilde{\psi}^{cap,l}$	Peg	Nonpeg
$\hat{\alpha}_1$	0.542 (11.50)	0.813 (11.95)	0.655 (9.11)	0.502 (7.28)	0.344 (3.55)	0.825 (14.69)	0.450 (8.16)
$\hat{\alpha}_2$	-0.025 (-1.06)	-0.051 (-1.00)	-0.111 (-1.63)	0.033 (1.41)	-0.072 (-1.55)	-0.134 (-2.11)	-0.024 (-0.91)
$\hat{\theta}_1$	0.622 (13.30)	0.836 (14.08)	0.731 (9.84)	0.564 (7.45)	0.459 (4.49)	0.847 (15.70)	0.535 (9.27)
$\hat{\theta}_2$	0.027 (1.12)	0.077 (2.13)	-0.071 (-1.21)	0.063 (2.93)	-0.006 (-0.12)	0.012 (0.23)	0.019 (0.75)

Parametric panel estimates for model  $\tilde{y}_{it} = \tilde{\mathbf{x}}'_{it}\alpha + \epsilon_{it}$  (upper part), as well as for the model  $y_{it}^* = \mathbf{x}_{it}^*\theta + \tilde{\epsilon}_{it}$  (lower part) where the impact of domestic policy rules is not taken into account. Transmission parameter estimates are provided for the full sample as well as for subsamples that are defined by capital mobility ( $\tilde{\psi}_{it}^{cap}$ ) and FX rate volatility ( $\tilde{\psi}_{it}^{fx}$ ) in country  $i$  and time  $t$ . The four different states are given by:  $\tilde{\psi}_{it}^{fx,h} \iff \tilde{\psi}_{it}^{fx} \geq \text{Med}(\tilde{\psi}_t^{fx})$ ;  $\tilde{\psi}_{it}^{fx,l} \iff \tilde{\psi}_{it}^{fx} < \text{Med}(\tilde{\psi}_t^{fx})$ ;  $\tilde{\psi}_{it}^{cap,h} \iff \tilde{\psi}_{it}^{cap} < \text{Med}(\tilde{\psi}_i^{cap})$  and  $\tilde{\psi}_{it}^{cap,l} \iff \tilde{\psi}_{it}^{cap} \geq \text{Med}(\tilde{\psi}_i^{cap})$ .  $t$ -statistics based on Newey-West standard errors are given in parentheses. Panel estimates according to the classification in Shambaugh (2004) are given in the right panel. The corresponding coefficient estimate  $\hat{\theta}_1$  for yearly data in Shambaugh (2004, p.325) is 0.79 for pegs and 0.55 for nonpegs under absence of capital controls.

Economies characterized by higher variation of FX rates and rather mobile capital ( $\tilde{\psi}^{fx,h}, \tilde{\psi}^{cap,h}$ ) are less subject to international transmission of interest rate changes ( $\hat{\alpha}_1 = 0.45$ ) while adjustment dynamics  $\hat{\alpha}_2$  towards the steady state equilibrium are insignificant. Lastly, economies with low capital mobility and rather flexible FX rates ( $\tilde{\psi}^{fx,h}, \tilde{\psi}^{cap,l}$ ) are least affected by interest rate transmission ( $\hat{\alpha}_1 = 0.34$ ).

Turning to estimates based on a similar classification as in Shambaugh (2004), pegs are characterized by high transmission of interest rate changes  $\hat{\alpha}_1 = 0.83$ , while also adjustment dynamics  $\hat{\alpha}_2 = -0.13$  are negative with 5% significance. In contrast, contemporaneous transmission of nonpegs' interest rate changes is weaker ( $\hat{\alpha}_1 = 0.45$ )

and adjustment dynamics are insignificant. The counterpart estimates for  $\hat{\alpha}_1$  in Shambaugh (2004, p.325) are 0.79 for pegs and 0.55 for nonpegs under absence of capital controls. However, comparing this evidence to our results is only justified if observed correlation of interest rates fully reflects monetary dependence. Therefore, it is of immediate interest if an exclusion of control variables  $\tilde{\pi}_{it-1}$ ,  $\tilde{y}_{it-1}$  and  $\tilde{s}_{ijt-1}$  from (4.1) yields (upward) biased estimates of monetary dependence especially for more flexible regimes under synchronized business cycles. To address this point, consider the restricted panel model

$$\begin{aligned} \Delta \xi_{it}^* &= \theta_1 \Delta \xi_{jt}^* + \theta_2 (\xi_{it-1} - \xi_{jt-1} - \phi_{ij,t-1}^{hp})^* + \tilde{e}_{it}, \\ \hat{=} \quad y_{it}^* &= \mathbf{x}_{it}^* \theta + \tilde{e}_{it}, \quad \forall \{i, t\} \in (\tilde{\psi}^{cap,\bullet}, \tilde{\psi}^{fx,\circ}), \end{aligned} \quad (4.5)$$

where e.g.  $y_{it}^* \equiv y_{it} - \bar{y}_i$  with  $\bar{y}_i$  denoting the sample mean of  $y_{it}$ . Table 4.2 provides regime specific estimates for panel model (4.5). Overall, ignoring the impact of domestic fundamentals appears to have two distinct effects. On the one hand interest rate transmission seems strongly upward biased for economies that are presumed most independent ( $\tilde{\psi}^{fx,h}$ ,  $\tilde{\psi}^{cap,l}$  or 'nonpegs'), since estimates  $\hat{\theta}_1$  exceed corresponding quantities  $\hat{\alpha}_1$  up to  $\hat{\theta}_1 - \hat{\alpha}_1 = 0.11$ . In turn, there is hardly any difference between  $\hat{\theta}_1$  and  $\hat{\alpha}_1$  for economies that should suffer from loss of monetary autonomy ( $\tilde{\psi}^{fx,l}$ ,  $\tilde{\psi}^{cap,h}$  or 'pegs'). These findings underline the argument that interest rates of (presumably) more independent economies should ceteris paribus incorporate more information about domestic fundamentals. Interestingly, estimated interest rate transmission obtained under absence of control variables for the group of nonpegs ( $\hat{\theta}_1 = 0.54$ ) is very close to the corresponding estimate of 0.55 in Shambaugh (2004). This might be a hint that the lack of control variables in former studies implies an understatement of independence especially for more flexible FX regimes. On the other hand, somewhat 'odd' parameter estimates  $\hat{\theta}_2$  are more reasonable after controlling for domestic fundamentals. As already suspected by Frankel et al. (2004) and Shambaugh (2004), our empirical findings support an economically significant share of common movements in interest rates to reflect synchronized business cycle behavior. Thus, evaluating monetary dependence merely in terms of interest rate comovement might not be fully appropriate. Moreover, the "Fear of Floating" phenomenon implying comovements in interest rates under de jure flexible FX rates could be less relevant if one adequately controls for domestic fundamentals and real economic linkages.

Unconditional (i.e. state invariant) estimates of interest rate response to domestic fundamentals obtained from the regression in (4.1) are given in table 4.3. In line with



the monetary independence of Germany and Japan that is postulated in Frankel et al. (2004), associated interest rates significantly respond to domestic gaps of output and inflation. Moreover, there is also evidence that interest rates of some smaller countries characterized by high FX flexibility (New Zealand) or even intermediate to low FX variation (Netherlands and Switzerland) might react to domestic fundamentals.

Table 4.3: Response to domestic fundamentals: Parameter estimates for single countries

	Aus	Aut	Be	Ca	Den	Fin	Fr	Ger	It	Ir
$\hat{\beta}_2$	0.120 (1.72)	-0.085 (-1.44)	0.071 (1.15)	0.009 (0.11)	0.048 (0.41)	-0.022 (-0.23)	0.098 (1.66)	0.337 (7.34)	-0.105 (-2.44)	0.019 (0.25)
$\hat{\beta}_3$	0.088 (1.71)	0.042 (1.66)	0.002 (0.33)	0.031 (0.98)	0.006 (0.57)	0.038 (1.16)	0.092 (1.90)	0.042 (5.23)	0.036 (1.56)	-0.006 (-0.51)
$\hat{\beta}_4$	-1.140 (-1.40)	-3.821 (-0.87)	1.141 (6.16)	-0.285 (-0.20)	3.627 (2.68)	-1.300 (-1.37)	3.602 (2.48)	0.318 (1.55)	0.007 (3.36)	-16.16 (-1.95)
	Jap	Ko	Nl	Nor	Nz	Sin	Sp	Swe	Swi	Uk
$\hat{\beta}_2$	0.196 (3.62)	-0.163 (-2.76)	-0.051 (-1.61)	-0.025 (-0.35)	0.142 (2.60)	0.079 (1.52)	0.286 (1.99)	-0.032 (-0.45)	0.072 (1.05)	0.197 (1.39)
$\hat{\beta}_3$	0.039 (1.44)	3E-4 (2.34)	0.014 (2.13)	-0.007 (-0.26)	0.000 (0.36)	0.000 (0.29)	0.004 (0.04)	0.016 (1.08)	0.124 (3.06)	0.015 (1.47)
$\hat{\beta}_4$	0.005 (1.29)	0.001 (2.41)	22.846 (1.60)	-0.170 (-0.32)	-0.904 (-1.39)	1.764 (1.59)	0.059 (1.19)	1.352 (3.78)	1.374 (0.70)	1.663 (0.74)

Parameter estimates for the response of interest rates to domestic fundamentals obtained from single country regressions according to model (4.1).  $t$ -statistics based on Newey-West standard errors are given in parentheses.

Moreover, countries with minor FX flexibility such as Belgium, Denmark, France, Italy and Sweden might have adjusted interest rates to stabilize their exchange rates. Accordingly, deviations from steady states of output and inflation do not have a significant impact on interest rate changes in these countries.

Concerning panel estimation of interest rate transmission, overly general classifications as used in this section might not be fully appropriate due to within group parameter heterogeneity. Moreover, such classifications are of limited applicability for policy advice since each economy might have its individual preference with regard to an acceptable degree of FX rate variation. The requirement for a thorough differentiation among states of FX flexibility motivates an assessment based on continuous functional coefficient panel models.

### 4.3 Functional coefficient panel models

In theory, the macroeconomic trilemma formalizes that the exposure of small economies' interest rates to world market rates should exhibit a continuous relation to its factual FX flexibility and the state of capital mobility. In the spirit of the functional model in Herwartz and Xu (2009) consider a semiparametric extension of (4.2)

$$\check{y}_{it} = \check{\mathbf{x}}'_{it}\alpha(\omega) + u_{it}, \quad \alpha(\omega) = (\alpha_1(\omega), \alpha_2(\omega)), \quad (4.6)$$

where  $\check{y}_{it}$ ,  $\check{\mathbf{x}}_{it}$  are obtained from the partial regression in (4.2). The specification in (4.6) is a local model for interest rate transmission dynamics in the sense that its parameters  $\alpha(\omega)$  depend on a representative continuum of states  $\omega$  that, for instance, summarize i) country specific exchange rate variability (discussed in section 4.3.1), or ii) both country specific exchange rate variability and capital mobility (section 4.3.2). Semiparametric functional coefficient estimators are subject to the curse of (factor) dimensionality. With respect to subsequent choices of factor variables, note that the factor dimension will not exceed 2 given the available sample size of quarterly observations. Technical details on estimation and inference in functional coefficient models are given in appendix A and appendix B, respectively.

#### 4.3.1 Dynamic state definitions: FX flexibility

For the moment assume that, during the last two decades, capital has been sufficiently and uniformly mobile such that the ability of small developed economies to influence domestic interest rates has been predominantly determined by the flexibility of FX rates. Let  $\psi_{it}^{fx}$  denote the logarithmic realized exchange rate volatility in economy  $i$  over the last quarter<sup>6</sup>, i.e.

$$\psi_{it}^{fx} = \ln(RV_{t-1}^{fx,ij}), \quad RV_t^{fx,ij} = \sum_{m \in t} (\Delta \ln s_{ij,m})^2,$$

where  $s_{ij,m}$  is the price of benchmark currency  $j$  in terms of domestic currency  $i$  in month  $m$ . In the following, all state variables (or factors, henceforth) are given in

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<sup>6</sup>Taking logs improves the distributional features of realized volatilities which tend to be strongly skewed to the left. Since, especially for the EMU members, FX volatilities may become zero, we add a small constant of 0.01 to all volatilities before taking logs.

standardized form to facilitate (cross sectional) overall comparability. Let

$$\omega_{it}^{(1)} = (\psi_{it}^{fx} - \bar{\psi}_i^{fx}) / \sigma_i(\psi_i^{fx}),$$

$$\text{with } \bar{\psi}_i^{fx} = \frac{1}{T} \sum_{t=1}^T \psi_{it}^{fx}, \quad \sigma_i(\psi_i^{fx}) = \sqrt{\frac{1}{T-1} \sum_{t=1}^T (\psi_{it}^{fx} - \bar{\psi}_i^{fx})^2}$$

be a measure of FX variability that characterizes the time path of (standardized) realized FX volatility for a given country. To contrast the state of FX flexibility for an economy in time  $t$  against other markets, define a standardized state variable

$$\omega_{it}^{(2)} = (\psi_{it}^{fx} - \bar{\psi}_t^{fx}) / \sigma_t(\psi_t^{fx}),$$

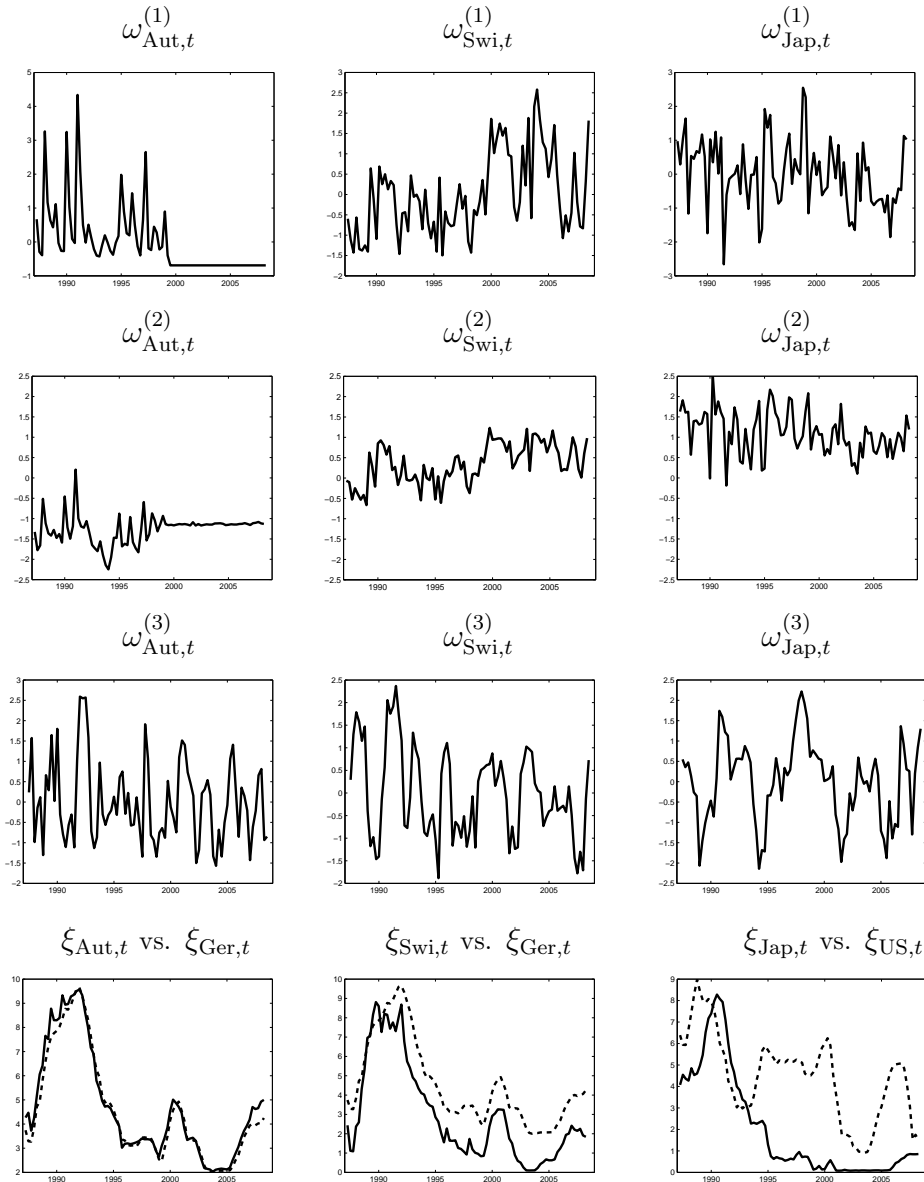
$$\text{with } \bar{\psi}_t^{fx} = \frac{1}{N} \sum_{i=1}^N \psi_{it}^{fx}, \quad \sigma_t(\psi_t^{fx}) = \sqrt{\frac{1}{N-1} \sum_{i=1}^N (\psi_{it}^{fx} - \bar{\psi}_t^{fx})^2}.$$

Finally, let  $\omega = (\omega^{(1)}, \omega^{(2)})'$  summarize the observable states of country specific exchange rate flexibility. For illustrational purposes, figure 4.1 sketches time paths of observed states  $\omega_{it}^{(1)}$  and  $\omega_{it}^{(2)}$  for Austria, Switzerland and Japan which may be seen to represent economies with unconditionally low, intermediate and high FX flexibility. Obviously, with respect to country specific variability over time and measured against other economies, we diagnose state dependence in both directions. For example, the Swiss Franc is characterized by an increase in FX rate volatility  $\omega_{it}^{(1)}$  after 1999. Moreover, this increase led to a reclassification (due to  $\omega_{it}^{(2)}$ ) in the sense that the Swiss Franc moved from a low FX volatility regime to more flexible states compared with other currencies. Notably, a high realization  $\omega_{it}^{(1)}$  does not necessarily imply flexible FX rates. It rather measures if an economy is characterized by a relatively high FX variation compared with its own historical experience. Accordingly, a high  $\omega_{it}^{(1)}$  in combination with a low  $\omega_{it}^{(2)}$  might reflect that a pegging economy has difficulties to maintain the peg.

Histograms of factor variables  $\omega_{it}^{(1)}$ ,  $\omega_{it}^{(2)}$  are shown in the upper panel of figure 4.2. The apparent peaks featuring FX based factors reflect quotes belonging to EMU members after the introduction of the Euro. Since intra EMU FX rates have been fixed since then, FX volatilities of all EMU countries were zero, implying a high concentration of observations in the corresponding area of the empirical support. Note that though the latter FX volatilities are zero, there are some observations falling below EMU implied state measures, which is due to the cross sectional standardization. In appendix C, additional empirical results are provided for a subsample of panel data

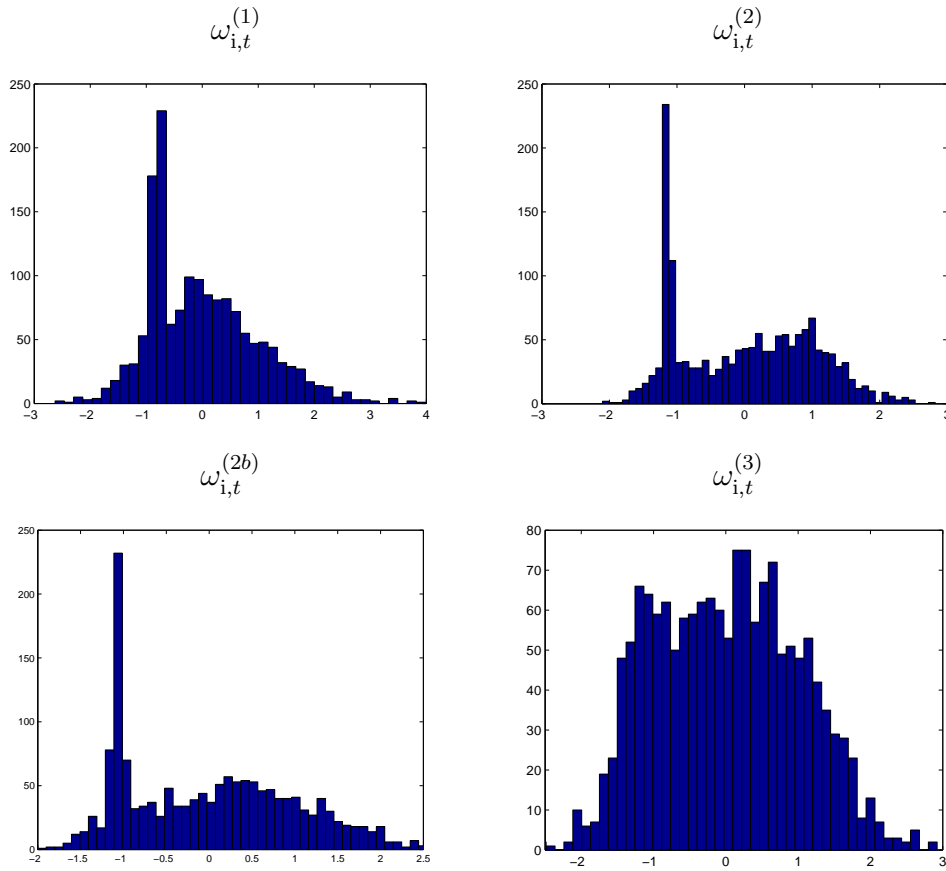
excluding quotes associated with intra EMU dynamics (i.e. interest rate transmission from Germany to Austria, Belgium, France, Italy, Netherlands, Spain, Finland and Ireland after 1998Q4) to show that these observations not decisive for our conclusions.

Figure 4.1: Selection of observed state variables and corresponding interest rates



Factor variables and interest rates including the corresponding reference rates (dashed) for a selection of economies (Austria, Switzerland and Japan). According to  $\omega_{it}^{(2)}$ , these economies might be (on average) considered representative for not flexible, intermediate and flexible currency regimes.

Figure 4.2: State variables: Empirical distribution



Histograms of factor variables measuring states of FX flexibility (i.e.  $\omega_{i,t}^{(1)}$ ,  $\omega_{i,t}^{(2)}$ ,  $\omega_{i,t}^{(2b)}$ ) and capital mobility ( $\omega_{i,t}^{(3)}$ ). Formal definitions are given in section 4.3.1, section 4.3.2 and appendix B.

### 4.3.2 Dynamic state definitions: The macroeconomic trilemma

The macroeconomic trilemma suggests that monetary autonomy is lost if capital is fully mobile and exchange rates are stable. Hence, in addition to measures of FX rate stability, (country specific) measures of capital mobility should be taken into account explicitly as determinants of monetary dependence. We rely on  $\omega_{it}^{(2)}$  (i.e. cross-sectionally evaluated FX volatility) for measuring FX variation. Capital mobility and capital market globalization has often been approximated by the extent of international real interest rate equalization (e.g. Mishkin 1984, Obstfeld and Taylor 2002). Let  $\psi_{it}^{cap}$  denote the logarithmic<sup>7</sup> absolute real interest rate differential between economy  $i$  and benchmark country  $j$  prevailing in time  $t - 1$ . Formally,  $\psi_{it}^{cap} = \ln(|r_{it-1} - r_{jt-1}|)$ .

<sup>7</sup>Similar to section 4.3.1, we add a small constant and take logs to reduce the skewness of the factor distribution.

We define real interest rates  $r_{it}$  as before in section 2.2. Moreover, real interest rate differentials enter in lagged form to avoid endogeneity problems. Summarizing country specific time paths of capital mobility over the last 20 years, a state measure for capital mobility is defined as

$$\omega_{it}^{(3)} = (\psi_{it}^{cap} - \bar{\psi}_i^{cap}) / \sigma_i(\psi_i^{cap}),$$

with  $\bar{\psi}_i^{cap} = \frac{1}{T} \sum_{t=1}^T \psi_{it}^{cap}$ ,  $\sigma_i(\psi_i^{cap}) = \sqrt{\frac{1}{T-1} \sum_{t=1}^T (\psi_{it}^{cap} - \bar{\psi}_i^{cap})^2}$ .

To underline that our results are robust to alternative measures of capital mobility, we additionally provide functional estimates based upon the (time varying) Chinn-Ito financial openness index (Chinn and Ito 2008) in the appendix. Finally, let  $\omega = (\omega^{(2)}, \omega^{(3)})'$  summarize states of exchange rate stability and country specific capital market integration. Dynamic and unconditional characteristics of the estimated states of capital market integration are displayed in figures 4.1 and 4.2, respectively.

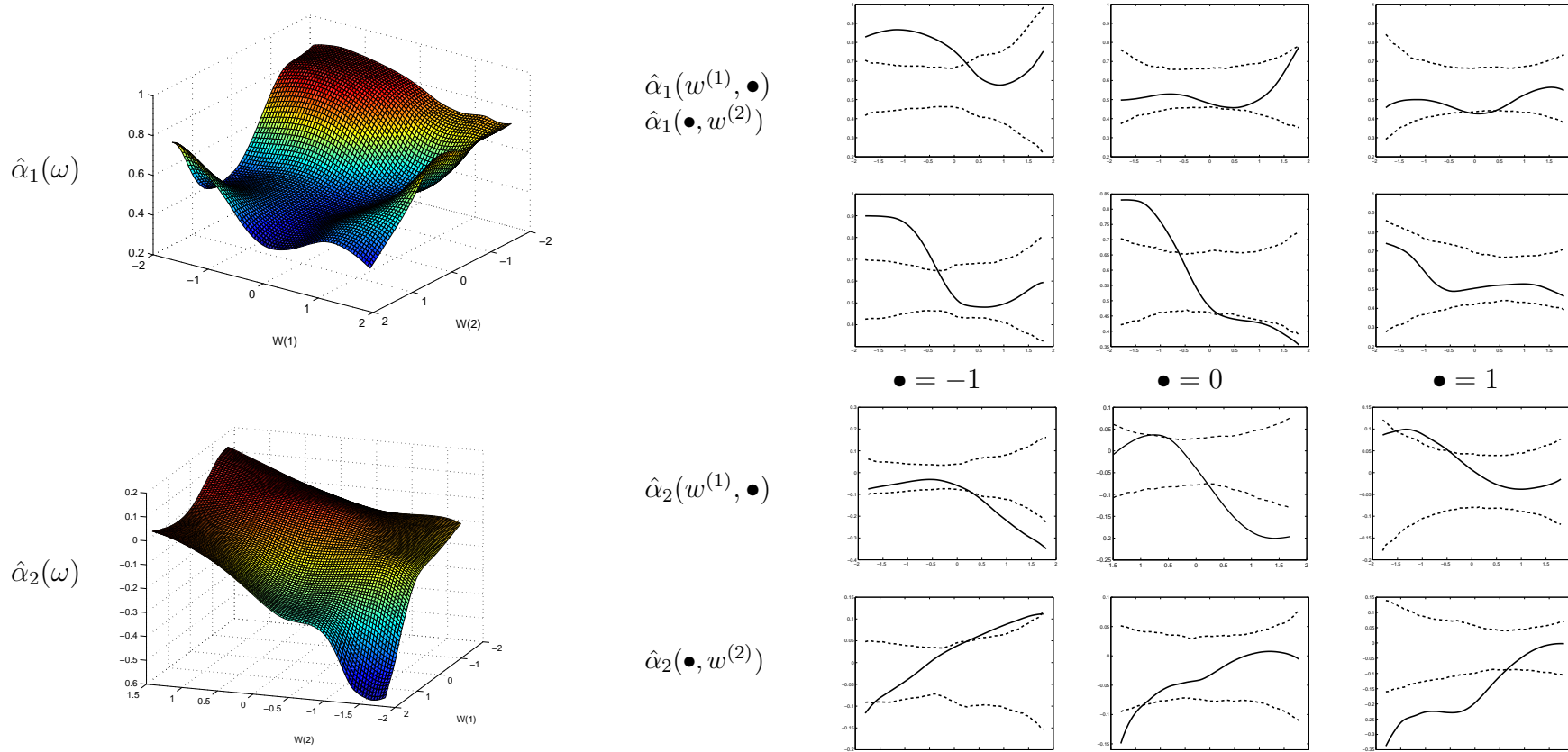
### 4.3.3 Results

#### Monetary independence as a function of FX rate flexibility

Figure 4.3 displays parameter estimates for the two-factor functional coefficient model obtained over a representative continuum of states measuring exchange rate flexibility defined in section 4.3.1. At first, one observes that, quantified by  $\hat{\alpha}_1(\omega)$  and  $\hat{\alpha}_2(\omega)$ , international transmission decreases over states of relatively high FX variation. Accordingly, results for global tests for factor invariance of transmission dynamics given in table 4.4 suggest this functional relationship to be globally significant.

Locally, figure 4.3 shows that economies which are characterized by a low realized FX volatility in comparison with other economies suffer from the highest contemporaneous transmission  $\hat{\alpha}_1(\omega)$  of reference nominal interest rates. Given that a currency is in a state of relatively low country specific FX variability (i.e.  $\omega^{(1)} \approx -1$ ), and, moreover, this variation is rather small with respect to the other currencies in the sample (e.g.  $\omega^{(2)} \approx -1$ ), monthly contemporaneous interest rate transmission  $\hat{\alpha}_1(\omega) = 0.9$  is close to unity, implying almost full transmission of world interest rates.

Figure 4.3: Functional coefficient estimates  $\hat{\alpha}_1(\omega), \hat{\alpha}_2(\omega)$ : FX flexibility



The upper (lower) panel presents a three dimensional plot of functional coefficient  $\hat{\alpha}_1(\omega)$  ( $\hat{\alpha}_2(\omega)$ ), where  $\omega = (w^{(1)}, w^{(2)})$  and corresponding selected partial (i.e. two dimensional) functional relations with confidence intervals given in dashed lines. For instance, the first row in the upper right panel shows selected relationships  $\hat{\alpha}_1(w^{(1)}, -1)$ ,  $\hat{\alpha}_1(w^{(1)}, 0)$  and  $\hat{\alpha}_1(w^{(1)}, 1)$ , while the second row presents  $\hat{\alpha}_1(-1, w^{(2)})$ ,  $\hat{\alpha}_1(0, w^{(2)})$  and  $\hat{\alpha}_1(1, w^{(2)})$ .

Table 4.4: Factor dependence: Global inference

	$(\omega^{(1)}, \omega^{(2)})$	$(\omega^{(2)}, \omega^{(3)})$	$(\omega^{(2b)}, \omega^{(3)})$	$(\omega^{(2)}, \omega^{(3b)})$	$(\omega^{(2)}, \omega^{(3)})^b$
$\hat{u}'\hat{u}/TN$	0.216	0.211	0.214	0.217	0.263
$(\omega^{(p*)}, \omega^{(q)})$	.022	.006	.090	.004	.069
$(\omega^{(p)}, \omega^{(q*)})$	.002	.000	.001	.000	.038
$(\omega^{(p*)}, \omega^{(q*)})$	.000	.000	.000	.000	.006

Test for global factor dependence based on the fit  $\hat{u}'\hat{u}/TN$  of model (4.6).  $p$ -values obtained by resampling according to appendix A are given in the lower three rows. Resampling is denoted with asterisks. For instance,  $(\omega^{(p*)}, \omega^{(q)})$  implies that only the first factor is resampled, while  $(\omega^{(p*)}, \omega^{(q*)})$  indicates a 2-tupel wise resampling from  $(\omega^{(p)}, \omega^{(q)})$ . Accordingly, these tests account for partial and joint factor invariance of interest rate transmission, respectively.

In turn, economies with relatively high realized FX volatility ( $\omega^{(2)} \approx 1$ ) are less severely subject to contemporaneous interest rate transmission varying between  $\hat{\alpha}_1(\omega) = 0.4$  and  $\hat{\alpha}_1(\omega) = 0.5$ . Since corresponding local confidence intervals do not include zero transmission  $\alpha_1(\omega) = 0$ , we reject interest rate autonomy for small economies with rather flexible FX rates. Generally, state dependence of  $\hat{\alpha}_1(\omega)$  is locally significant, since estimates often go beyond bootstrap based 95% confidence intervals which are derived under the assumption of state invariance (see appendix A). For example, given that country specific FX variation is close to its historic average ( $\omega^{(1)} \approx 0$ ), we observe transmission parameter estimates below confidence intervals if corresponding FX variation exceeds the global mean ( $\omega^{(2)} \geq 0$ ). In turn, functional transmission estimates are above confidence bounds if observed FX variation falls short of  $\omega^{(2)} \leq -0.5$ .

Measuring the adjustment speed towards a steady state nominal interest rate differential, we observe two state specific characteristics for  $\hat{\alpha}_2(\omega)$ . Firstly, nominal interest rates adjust (significantly) faster when exchange rate volatility is relatively low in comparison with the cross sectional average level of FX rate volatility ( $\omega^{(2)} \approx -1$ ). Moreover, if the domestic currency features higher variability than most other currencies (e.g.  $\omega^{(2)} \approx 1$ ), parameter estimates are close to or even above zero which means that domestic rates do not adjust to the steady state. In line with adjustment parameter estimates from conventional parametric panel models in section 4.2.4, state invariance implied local confidence intervals include  $\hat{\alpha}_1(\omega) = 0$ . In contrast, highlighting the merits of local estimation, state dependent estimates are often not included in the interval. For all levels of country specific variation, the speed of adjustment significantly increases in the cross sectional measure of FX stability  $\omega^{(2)}$ . For low values of  $\omega^{(2)}$ ,  $\hat{\alpha}_2(\omega)$  approaches  $-0.2$ , implying a half life of deviations from steady



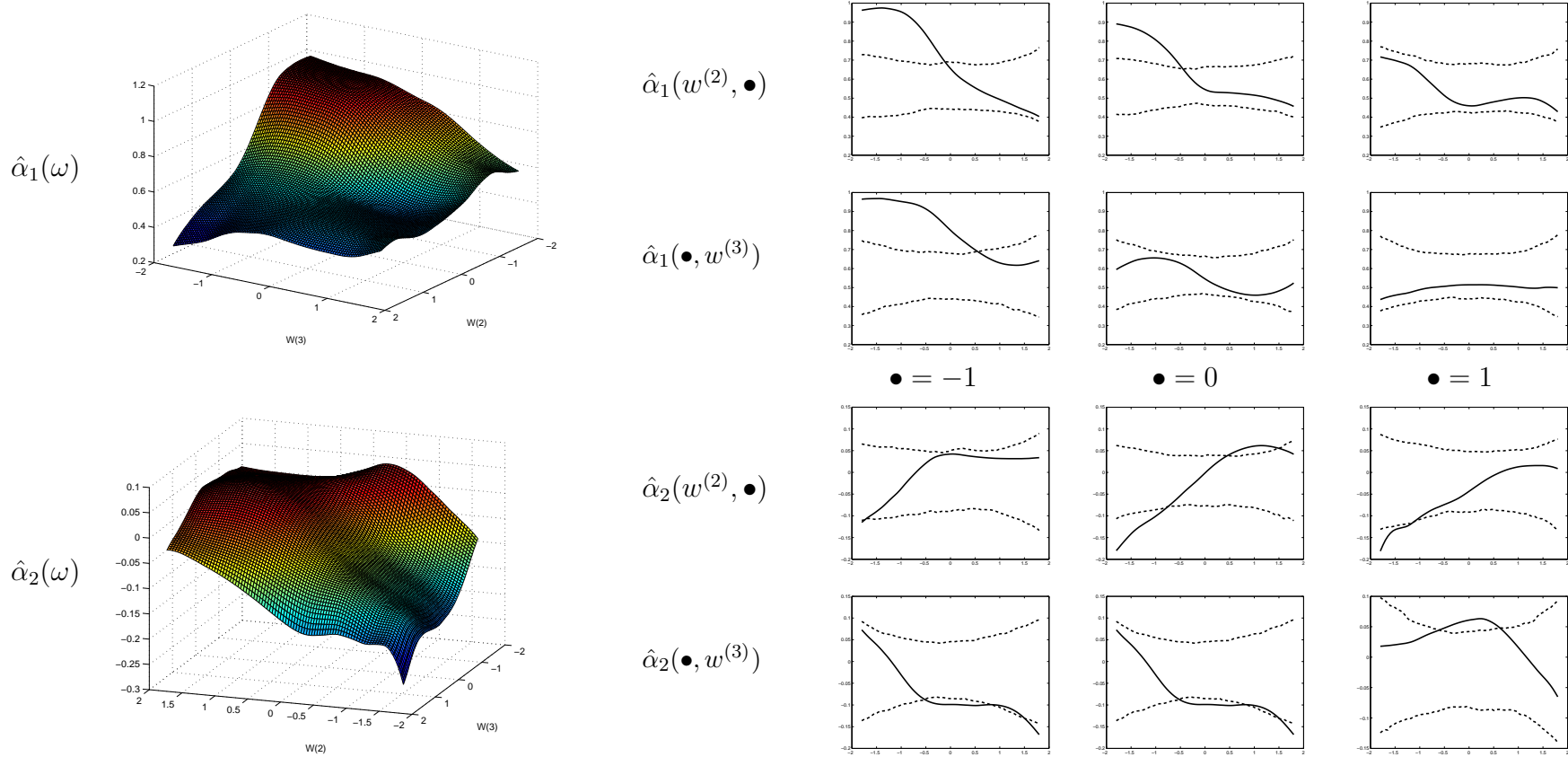
state interest differentials of approximately three quarters. Secondly, we observe that the adjustment speed also significantly increases in  $\omega^{(1)}$ , with fastest estimated adjustment diagnosed conditional on states of high country specific volatilities  $\omega^{(1)}$  and relatively low states  $\omega^{(2)}$ . This observation could reflect scenarios in which the stability of (pegged) currencies is threatened by speculative market forces. Monetary authorities might then rise interest rates so markedly that speculators refrain from going short in the domestic currency (Obstfeld and Rogoff, 1995). Such (short lived) peaks might tend to dominate the estimated adjustment of interest rates locally for states characterized by  $\{\omega^{(1)} > 1, \omega^{(2)} < -1\}$ .

### Monetary independence implied by the macroeconomic trilemma

Figure 4.4 shows parameter estimates for the two-factor functional coefficient model obtained over different (i.e. representative) states summarizing FX variability ( $\omega^{(2)}$ ) and capital market integration ( $\omega^{(3)}$ ) as defined in section 4.3.2. Refining the overall evidence from parametric panel based evidence in section 4.2.4, functional estimates  $\hat{\alpha}_1(\omega)$  and  $\hat{\alpha}_2(\omega)$  suggest a marked trade-off among FX stability, capital mobility and monetary independence. Hence, the implications of the macroeconomic trilemma seem to hold for developed economies over the last two decades.

Confirming this overall impression formally, we reject global factor invariance of transmission dynamics at conventional significance levels according to test results in table 4.4. Moreover, observed transmission depends on both capital mobility and FX flexibility since partial factor invariance is rejected throughout. Thus, the extent of FX flexibility and capital market integration matters for interest rate transmission in general. Locally, figure 4.4 shows estimated contemporaneous interest rate transmission  $\hat{\alpha}_1(\omega)$  to increase with exchange rate stability and capital market integration (note that high values of  $\omega^{(3)}$  indicate states of relative market disintegration). States of rather low FX rate volatility and high capital mobility are characterized by contemporaneous transmission estimates  $\hat{\alpha}_1(\omega)$  close to unity that steadily decay to  $\hat{\alpha}_1(\omega) = 0.6$  if capital mobility decreases (*ceteris paribus*).

Figure 4.4: Functional coefficient estimates  $\hat{\alpha}_1(\omega)$ ,  $\hat{\alpha}_2(\omega)$ : The Trilemma



The upper (lower) panel presents a three dimensional plot of the functional coefficient estimate  $\hat{\alpha}_1(\omega)$  ( $\hat{\alpha}_2(\omega)$ ), where  $\omega = (w^{(2)}, w^{(3)})$  and corresponding selected partial (i.e. two dimensional) functional relations with confidence intervals given in dashed lines. For instance, the first row in the upper right panel shows selected relationships  $\hat{\alpha}_1(w^{(2)}, -1)$ ,  $\hat{\alpha}_1(w^{(2)}, 0)$  and  $\hat{\alpha}_1(w^{(2)}, 1)$ , while the second row presents  $\hat{\alpha}_1(-1, w^{(3)})$ ,  $\hat{\alpha}_1(0, w^{(3)})$  and  $\hat{\alpha}_1(1, w^{(3)})$ .

In turn, if FX flexibility (i.e.  $\omega^{(2)}$ ) increases, estimated transmission quickly drops to values close to  $\hat{\alpha}_1(\omega) = 0.4$ . Overall, it appears that a higher share of coefficient variation is caused by FX flexibility in comparison with capital mobility induced variation. This seems intuitive since free flow of capital had been already established for the period under investigation. Local confidence intervals underline that states of high FX flexibility (low capital mobility) are characterized by local state invariance of capital mobility (FX flexibility). Hence, under relatively immobile capital (flexible FX rates), interest rate transmission is not subject to FX stability (capital mobility) any longer. However, confidence bands never include  $\hat{\alpha}_1(\omega) = 0$  and thus, full independence might not prevail for small economies even under high FX flexibility and/or low capital mobility. Accordingly, the predictions of the macroeconomic trilemma most likely hold in a 'relative' form. Similar to the evidence in figure 3, it seems that interest rate transmission rises if  $\omega^{(2)}$  falls below a certain threshold value  $\omega^{(2)} \approx 0$ . Therefore, an economy with high FX flexibility might stabilize its currency to a certain extent without suffering from an increasing exposure to changes in foreign interest rates. Functional estimates  $\hat{\alpha}_2(\omega)$  reveal that the adjustment of domestic interest rates to their steady state differential tends to be the faster the less flexible are corresponding FX rates. Again, parameter variation seems to be predominantly governed by states of FX flexibility  $\omega^{(2)}$ , while (in line with evidence based on parametric models) we also diagnose higher adjustment speed in states of lower capital market integration. Since market integration is measured by means of the real interest rate differential, relatively high capital market disintegration coupled with relatively low FX variability implies raised currency risk premia for a given economy. To some extent, this state (and corresponding local adjustment parameter estimates) should coincide with the case where (target) FX rates are threatened by speculative market forces discussed in section 4.3.3. In contrast to the local behavior of  $\hat{\alpha}_1(\omega)$ , we find that the foreign influence on domestic rates in terms of  $\hat{\alpha}_2(\omega)$  increases whenever FX rate volatility decreases. Thus, even stabilizing the domestic currency moderately might be accompanied by a stronger adjustment to the steady state interest rate differential. Also reflecting empirical results from section 4.3.3, we conclude that free floating economies might moderately attenuate FX rate movements without substantially increasing their exposure to foreign interest rates. However, economies with average FX rate flexibility substantially forfeit remaining degrees of freedom of domestic monetary policy by further decreasing flexibility.

## **4.4 Conclusions**

The macroeconomic trilemma suggests a binding trade-off among three fundamental aims of monetary authorities: To benefit from full capital mobility, to minimize variation in FX rates and to be able to conduct an independent monetary policy targeting at domestic goals. Since presumed gains in monetary autonomy are the main motivation to accept some variation in FX rates, a comprehensive assessment of this trade-off is a core prerequisite for monetary policymakers' decisions on the appropriate extent of FX rate flexibility. In this study we put particular effort on providing an extensive empirical assessment of the macroeconomic trilemma and its implications for monetary independence of developed economies by means of a flexible and uniform econometric panel framework.

Evidence from conventional parametric panel models indicates that, in line with Shambaugh (2004) and in contrast to Frankel et al. (2004), interest rates of more flexible currency regimes are less subject to exposure to foreign interest rates. Moreover, we argue that former evidence might have suffered from biased transmission estimates due to the lack of control variables ruling out the effects of synchronized business cycle behavior. Flexible semiparametric panel model estimates and corresponding tests strongly support the general validity of the macroeconomic trilemma, where transmission seems to be a positive (nonlinear) function of FX rate stability and capital mobility. Countries with fixed FX rates and closely integrated capital markets share one-to-one interest rate movements with their reference economies, while states that are either characterized by higher FX variation or lower capital mobility (or both) feature markedly less transmission. However, as interest rate transmission does not appear to linearly increase in FX rate flexibility and capital mobility, there is some scope for attenuating FX rate fluctuations at little cost, since (at least for more flexible regimes) marginal decreases in FX rate flexibility do not necessarily imply a marked strengthening of transmission. Accordingly, gains in monetary independence are substantial if a country abdicates pegging its FX rate, but the marginal benefit of tolerating higher FX flexibility quickly vanishes. Lastly, this study focuses on monetary dependence of developed economies. Besides data availability problems, developing economies are not considered since they deserve a different (separate) treatment with respect to their excessively volatile interest rate characteristics, risk premia and choice of domestic fundamentals. Studying determinants of monetary autonomy for developing countries in a modified (semiparametric) model is considered as an issue for future research.

# Appendix A

## Estimation and Inference

Estimation of the functional coefficient vector  $\alpha(\omega)$  is implemented as a multivariate version of the Nadaraya Watson estimator (Nadaraya, 1964; Watson, 1964). The estimator is defined as  $\hat{\alpha}(\omega) = \mathcal{X}^{-1}(\omega)\mathcal{Y}(\omega)$  and may be seen as a pooled weighted least squares estimator where observations  $\omega_{it}$  close to  $\omega$  enter with higher weights than observations  $\omega_{it}$  deviating more from  $\omega$ . The latter weighting procedure is implemented via multiplicative kernel functions. Formally,

$$\begin{aligned}\mathcal{X}(w^{(p)}, w^{(q)}) &= \sum_{i=1}^N \sum_{t=1}^T \tilde{\mathbf{x}}_{it} \tilde{\mathbf{x}}'_{it} K_h(\omega_{it}^{(p)} - \omega^{(p)}) K_h(\omega_{it}^{(q)} - \omega^{(q)}), \\ \mathcal{Y}(w^{(p)}, w^{(q)}) &= \sum_{i=1}^N \sum_{t=1}^T \tilde{\mathbf{x}}_{it} y_{it} K_h(\omega_{it}^{(p)} - \omega^{(p)}) K_h(\omega_{it}^{(q)} - \omega^{(q)}),\end{aligned}$$

where  $\{(p, q)\} \in \{(1, 2); (2, 3)\}$ . Moreover,  $K_h(\bullet) = \frac{1}{h} K(\frac{\bullet}{h})$  and  $K$  denotes a quartic kernel function, i.e.  $K(u) = \frac{15}{16}(1 - u^2)^2 I(|u| \leq 1)$ . For bandwidth selection, we use Scott's rule of thumb (Scott 1992), multiplied by a factor of 1.8. Smaller bandwidths turn out to eventually result in numerical problems that occur in regions of very sparse data involving singularity of local regression design matrices.

Local inference on  $\alpha(\omega)$  is based on bootstrap-based confidence intervals for parameters associated with international interest rate transmission  $\alpha_1(\omega), \alpha_2(\omega)$ . To infer if functional coefficient estimates are locally state dependent, the following factor based resampling scheme has been proposed by Herwartz and Xu (2009):

1. Local parameter estimates for model (4.6) can be considered as a function of the data and the chosen bandwidth parameter, i.e.

$$\hat{\alpha}_i(\omega) = f(y_{it}, \tilde{\mathbf{x}}'_{it}, \omega_{it} = (\omega_{it}^{(p)}, \omega_{it}^{(q)}), h, i = 1, \dots, N, t = 1, \dots, T). \quad (4.7)$$

2. Factor dependence and invariance of coefficients in model (4.7) are distinguished by a comparison among local coefficient estimates and corresponding bootstrap coefficient estimates

$$\hat{\alpha}^*(\omega) = f(y_{it}, \tilde{\mathbf{x}}'_{it}, \omega_{it}^* = (\omega_{it}^{(p)*}, \omega_{it}^{(q)*}), h, i = 1, \dots, N, t = 1, \dots, T). \quad (4.8)$$

Bivariate tuples  $\omega_{it}^* = (\omega_{it}^{(p)*}, \omega_{it}^{(q)*})$  are drawn with replacement from the set of

bivariate variables  $\{\{\omega_{it}^{(p)}, \omega_{it}^{(q)}\}_{i=1}^N\}_{t=1}^T$ . Note that sample information on  $y_{it}, \check{\mathbf{x}}'_{it}$  is not affected by the bootstrap. Therefore, the proposed scheme will generate a factor variable that, per construction, is independent of the functional coefficients. If the null hypothesis of state invariance is true, estimates  $\hat{\alpha}(\omega)$  and  $\hat{\alpha}^*(\omega)$  should only marginally deviate from each other if evaluated over the support of the factor variable.

3. A large number of draws,  $R = 1000$  say, of bootstrap estimates  $\hat{\alpha}^*(\omega)$  is considered as sufficient to approximate the underlying distribution under the null hypothesis of state invariance. For inferential purposes, estimates  $\hat{\alpha}(\omega)$  are shown with confidence intervals that present the 25th and 975th order statistic of  $\hat{\alpha}^*(\omega)$ . In this sense, the actual estimate is regarded to differ locally from the unconditional relation with 5% significance if the local confidence interval does not include the local estimate  $\hat{\alpha}(\omega)$ .

Local inference by means of confidence intervals is informative to characterize functional relationships conditional on specific economic states  $\omega$ . The same factor based resampling scheme applies to test for 'overall' factor dependence of interest rate transmission  $\alpha_1(\omega)$  and long term adjustment  $\alpha_2(\omega)$ . For this purpose we consider the residual sum of squares (RSS) of the functional regression

$$\check{y}_{it} = \check{\mathbf{x}}'_{it}\alpha(\omega) + u_{it}.$$

The null hypothesis of overall factor invariance,  $H_0 : \alpha(\omega) = \alpha$  is rejected with 5% significance if the sample based RSS statistic exceeds the 95% quantile of the respective distribution of bootstrap based counterparts. Notably, the introduced factor resampling scheme applies for testing the hypothesis of global joint factor invariance. To test for partial factor invariance, we draw with replacement from one factor only, while letting the other factor (and thus its potential link to the data) unchanged.

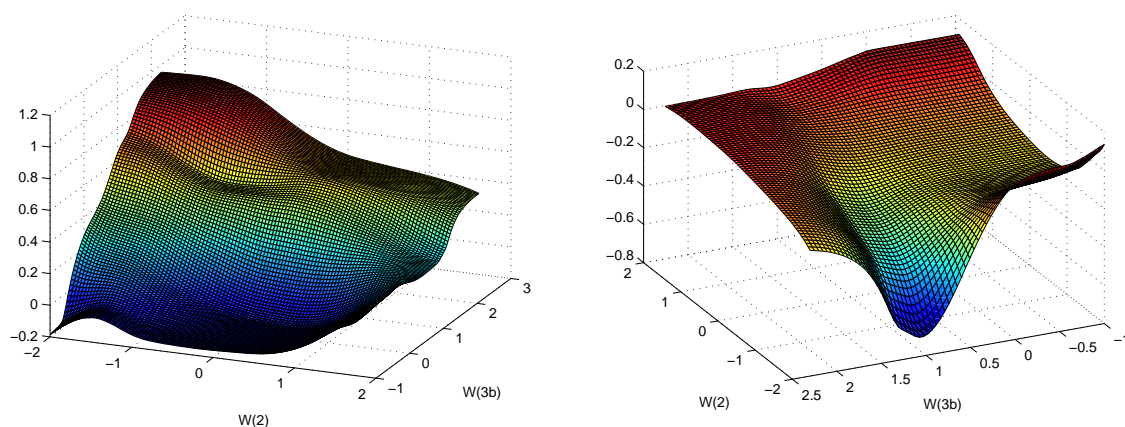
## Appendix B

### Robustness of empirical results

#### B.1 Using the Chinn-Ito index of financial openness

To highlight the robustness of our results, we further estimate functional interest rate transmission conditional on exchange rate stability  $\omega^{(2)}$  and the time varying Chinn and Ito index of financial openness, denoted as  $\omega^{(3b)}$  (Chinn and Ito 2008). The index does not enter in standardized form since it is comparable over distinct countries and its empirical range is very similar to the range of the other factors. Interest rate transmission estimates conditional on  $(\omega^{(2)}, \omega^{(3b)})$  are given in the upper panel of figure 4.5. At first sight, the estimated functional patterns in figure 4.5 appear very similar to those presented in figure 4.4. Contemporaneous transmission is literally zero in states of low capital mobility. Moreover, states of high FX variability are also characterized by very weak transmission. Given that capital mobility and/or FX stability increase, contemporaneous transmission increases. Global test results in table 4.4 reveal that the functional pattern is highly significant. Thus, our main findings appear robust with respect to alternative measures of capital mobility.

Figure 4.5: Robustness check I



Functional coefficient estimates evaluated conditional on  $\omega = (w^{(2)}, w^{(3b)})$ .

## B.2 Using an exchange rate band based factor variable

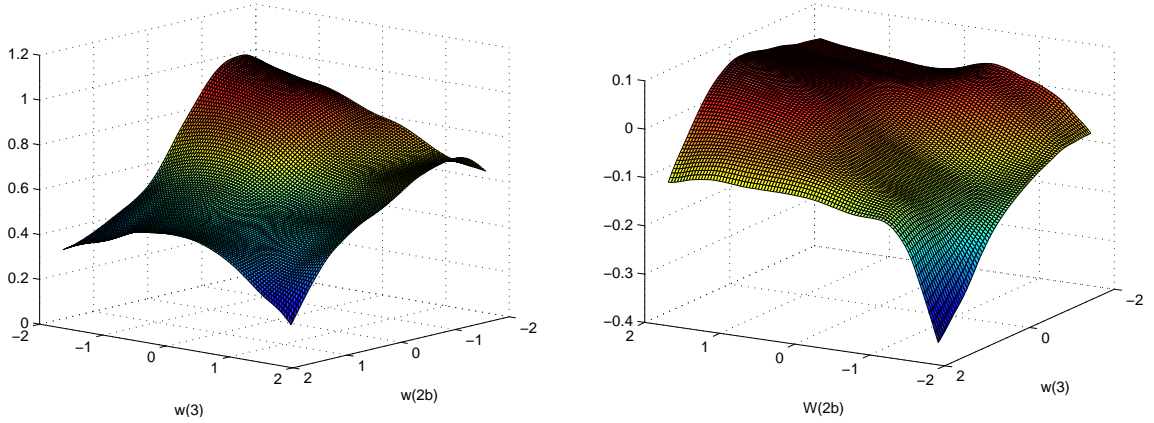
One natural way to generalize the exchange rate band approach (Shambaugh 2004) in a dynamic way is to consider the realized size of (hypothetical) FX bands. Define

$$\psi_{it}^{fxb} = \ln(\max(|\ln(s_{ij,m}) - \ln(\bar{s}_{ij,z})|)), \quad \forall m \in z; z = \{t-1, \dots, t-4\},$$

where  $\bar{s}_{ij,z}$  denotes the median FX rate (i.e. the presumed target value) that prevailed over the last twelve months. Moreover,  $\max(|\ln(s_{ij,m}) - \ln(\bar{s}_{ij,z})|)$  is the corresponding maximum percentaged absolute deviation from  $\bar{s}_{ij,z}$  which is considered in logs to reduce the skewness of the factor distribution. Furthermore,  $\omega_{it}^{(2b)} = (\psi_{it}^{fxb} - \bar{\psi}_t^{fxb})/\sigma_t(\psi_t^{fxb})$ , where  $\bar{\psi}_t^{fxb}$  is the empirical (time dependent) cross sectional mean of all country specific factors  $\bar{\psi}_t^{fxb} = \frac{1}{N} \sum_{i=1}^N \psi_{it}^{fxb}$ , and correspondingly,  $\sigma_t(\psi_t^{fxb}) = \sqrt{\frac{1}{N-1} \sum_{i=1}^N (\psi_{it}^{fxb} - \bar{\psi}_t^{fxb})^2}$  is their cross sectional dispersion.

Functional parameter estimates in the upper panel of figure 4.6 suggest that the overall result remains robust when substituting realized FX volatility with the realized width of a presumed FX band.

Figure 4.6: Robustness check II



Functional coefficient estimates conditional on  $\omega = (w^{(2b)}, w^{(3)})$ .

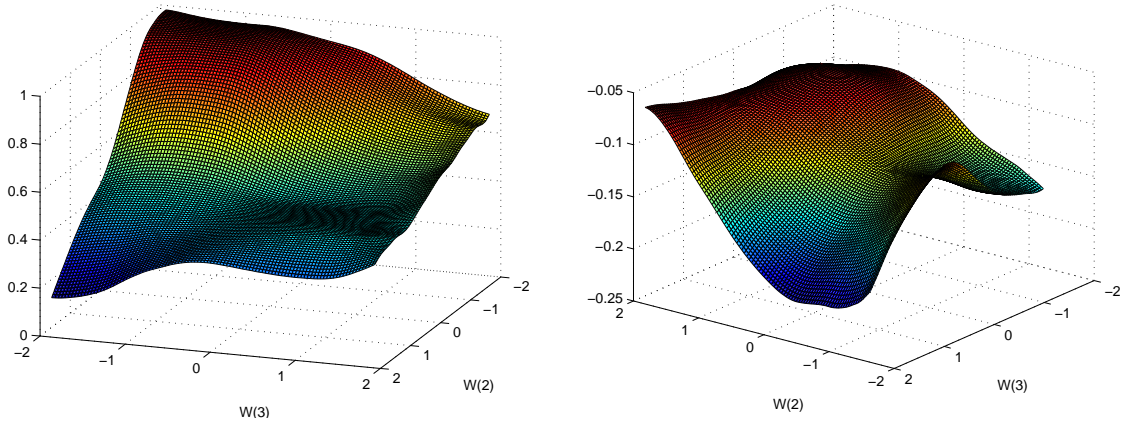
## B.3 Using interest rate differentials detrended in terms of OLS

We replace the 'pseudo error correction term',  $(\xi_{it-1} - \xi_{jt-1} - \phi_{ijt-1}^{hp})$ , in (4.6) by an interest rate differential detrended by means of OLS, namely  $(\xi_{it-1} - \xi_{jt-1} - \hat{\delta}_1 - \hat{\delta}_2 t - \hat{\delta}_3 t^2)$ . The augmentation with deterministic trend parameters ( $\hat{\delta}_2$  and  $\hat{\delta}_3$ ) allows for diminishing of time trends in presence of increasing market integration. Eyeballing figure 4.7 suggests that the functional adjustment pattern appears in line with the



benchmark case where HP filtered interest differentials are considered.

Figure 4.7: Robustness check III

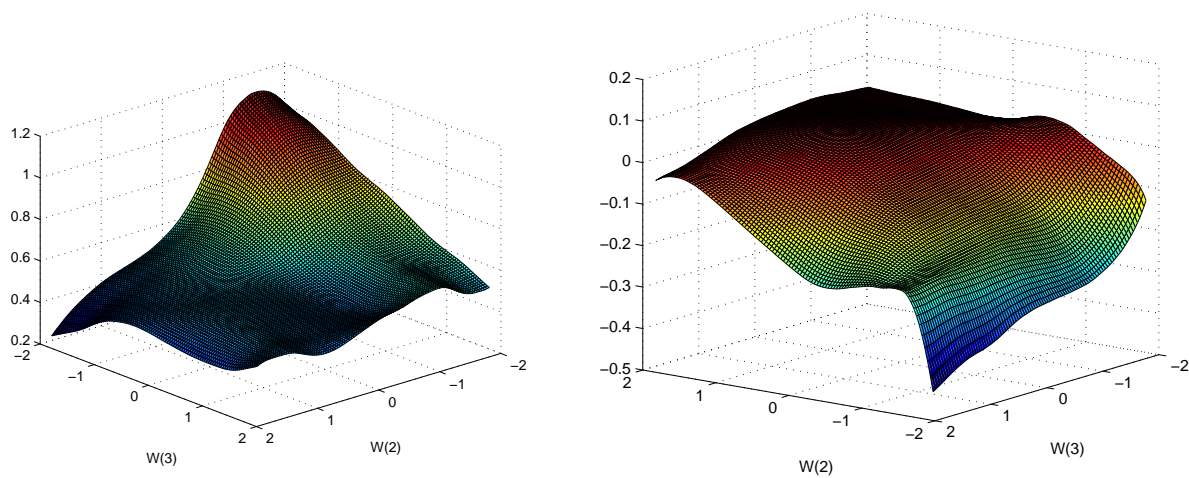


Functional coefficient estimates evaluated conditional on  $\omega = (w^{(2)}, w^{(3)})$ , where  $(\xi_{it-1} - \xi_{jt-1} - \phi_{ijt-1}^{hp})$  is replaced by the interest rate differential  $(\xi_{it-1} - \xi_{jt-1} - \hat{\delta}_1 - \hat{\delta}_2 t - \hat{\delta}_3 t^2)$  in model (4.6).

## B.4 Excluding intra EMU quotes

One might conjecture that observed functional characteristics in figure 4.4 just reflect two states implied by EMU and non-EMU observations. Hence, we additionally provide estimates for a subsample of panel data excluding intra EMU quotes (i.e. interest rate transmission from Germany to Austria, Belgium, France, Italy, Netherlands, Spain, Finland and Ireland after 1998Q4) to show that these observations are not decisive for our conclusions. Associated estimates conditional on  $(\omega^{(2)}, \omega^{(3)})^b$  excluding intra EMU quotes are given in figure 4.8. As a result, conclusions are qualitatively robust for the substantially smaller data set. Most importantly, local estimates that might be presumed to predominantly reflect observed transmission patterns within the EMU (i.e.  $\hat{\alpha}_1 \approx 1$  and  $\hat{\alpha}_2 \leq -0.2$  located in states of high capital mobility and low FX flexibility) hardly change. Accordingly, factor invariance is throughout rejected by means of respective diagnostics documented in table 4.4.

Figure 4.8: Robustness check IV



Functional coefficient estimates evaluated conditional on  $\omega = (w^{(2)}, w^{(3)})$  excluding intra-EMU quotes after 1998Q4.

## Chapter 5

# Monetary Independence II: The international linkage of the inflation climate

### 5.1 Introduction

A core concern of monetary policy is long term price stability. Nowadays, the framework of inflation targeting has become a widely followed strategy to implement a moderate and predictable evolution of prices. While the European Central Bank (ECB) considers inflation targeting as one of its two main monetary policy axioms, for instance, the Bank of Canada (BoC) or the Bank of England (BoE) explicitly follow this strategy. The so-called inflation forecast targeting provides a useful guideline for inflation targeting (Svensson 1997, Leitemo 2006). Inflation targeting and the corresponding adjustment of monetary instruments should be forward looking since expected inflation rates constitute a central determinant for a firms' setting of future prices. Since the essential element in such a framework is to anchor long run inflation expectations, the extent to which central banks are able to implement stable, uniform and definite beliefs is crucial. Bernanke (2003) notes that well-anchored inflation expectations facilitate the achievement of price stability in the long term and the stabilization of output and employment in the short run.

A central bank's ability to bring long term inflation in line with the announced target is an important aspect of the debate on currency regimes and monetary independence. The latter belongs to the lead arguments in favor of floating FX rates. Facing integrated markets, central banks of small open economies are subject to constraints that are summarized in the macroeconomic trilemma (Obstfeld, Shambaugh and Taylor 2005). The choice among a free flow of capital, a stable FX rate and inde-

pendent control over inflation targets is restricted, as only two of the three objectives can be achieved jointly. Central banks are considered to be able to set their long run domestic inflation rate independently at the 'cost' of floating FX rates. Empirical evidence, however, shows that flexible FX rates do not offer complete insulation. Building on an early study by Darby and Lothian (1989), Eun and Jeong (1999) and Yang, Guo and Wang (2006) examine international transmission of G7 inflation rates and detect strong linkages among these economies. Other studies address the issue of monetary independence by analyzing international linkage of nominal interest rates. Under market integration and floating FX rates, cross sectional independence of nominal interest rates hints at the ability to determine inflation expectations domestically. Frankel, Schmukler and Serven (2004) investigate the international transmission of nominal interest rates for different currency regimes and document that both floating and fixed FX rate regimes show transmission dynamics, with slower adjustments featuring the former. According to Frankel et al. (2004) the US, Japan and Germany are the only economies capable to set their own interest rates.

While the empirical literature documents nominal linkages, it is rarely discussed why nominal linkages prevail under floating FX regimes. Calvo and Reinhart (2002) point out that 'a pure float is an artifact of economics textbooks' and justify nominal linkages under flexible FX regimes with a 'fear of floating'. The latter notion refers to occasional FX market interventions likely motivated by strong pass through from FX rates to prices or liabilities in other currencies (Calvo and Reinhart 2002, Hausmann, Panizza and Stein 2001). Confirming this view, recent evidence by Lubik and Schorfheide (2007) underpins the relevance of (trade weighted) nominal exchange rates in policy rules presumed for the BoC or BoE. With respect to existing trade relations, long term inflation of small economies such as the UK (Canada) should especially be subject to the monetary policy in the EMU/US (US). A closer investigation of the determinants of nominal linkages is of immediate interest for monetary authorities and open macro economists.

In this chapter we empirically assess first and second order linkages of inflation compensation as a measure of nominal dependence in a free floating subsystem of the G7 comprising Canada (CA), France (FR), the UK and the US. In particular, our interest is to clarify if these linkages could be traced back to the influence of monetary policies of major central banks (such as the ECB or the US Federal Reserve System (Fed)) on the smaller economies. Moreover, the role of imperfect FX rate flexibility as a potential reason for nominal linkages is highlighted.

We contribute to the existing literature in several directions concerning the data,

the measure of dependence and the applied econometric methodology. While recent assessments of monetary independence (Eun and Jeong 1999, Yang et al. 2006, Frankel et al. 2004, Obstfeld et al. 2005) rely on inflation or nominal interest rate linkages, we investigate the degree to which ex-ante inflation rates are determined on international markets. We use inflation compensation as implied by the (liquidity adjusted) difference between yields of 'Treasury Inflation Protected Securities' (TIPS, hereafter) and corresponding conventional bonds. So-called breakeven inflation rates are regarded to measure future inflation prospects. For an accurate measurement of nominal dependencies, the use of breakeven inflation rates offers important advantages. Compared with short term interest rates comovements of long term inflation compensation should less reflect synchronized real business cycle dynamics triggering common dynamics of national policy interest rates. Moreover, linkages of nominal interest rates might reflect cross market features of both real rates or inflation compensation. Compared with inflation rates breakeven rates instantaneously summarize the available information about the nominal effects of current and future monetary policy. Notably, quantifying nominal dependence by means of realized inflation rates might be criticized since, besides inflation expectations, also past inflation rates and the current output gap determine current inflation rates.

In contrast to former studies, we provide evidence based on both first and second order characteristics. A vector autoregressive model (VAR) is employed to assess transmission (Granger causality) among first order moments. Moreover, second order characteristics are quantified by means of realized correlations. Finally, the comovements among inflation compensations are related to observable measures of *i*) global shocks, *ii*) monetary policies of dominating economies (i.e. EMU and the US) and *iii*) country specific FX rate flexibility.

To preview the results, VAR estimates reveal multidirectional patterns of Granger causality among (mostly nonstationary) international breakeven inflation rates. Empirical correlations among breakeven rates differ from zero, with stronger realized correlations diagnosed for closely linked markets. In particular long run inflation prospects in CA vary with expectations concerning long run US inflation and, thus, fail to fully reflect the domestic inflation target. Similarly, also inflation expectations in the Euro area<sup>1</sup> and the UK are subject to common shocks, but empirical correlations vary on a lower scale. Thus, in comparison with the BoC, the BoE is likely better able to target domestic long run inflation. Taking a dynamic perspective, em-

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<sup>1</sup>The French breakeven rate is assumed to represent inflation expectations in the Euro area. The ECB also refers to the French breakeven inflation rate in its Monthly Bulletins.

empirical findings suggest that Canadian and US rate correlations are relatively high if *i*) the US conducts an expansionary monetary policy, *ii*) US inflation uncertainty is high or *iii*) if the CA/US FX rate volatility is low. Similarly, UK and EMU linkage is strengthened if (long term) inflation expectations in the Euro area become relatively uncertain. Nominal linkage between the UK and the US rises with US inflation uncertainty and strong fluctuations in UK FX reserves that likely indicate FX market interventions of the BoE.

In summary, we find evidence that inflation prospects in small economies with free floating currencies are influenced by monetary policy related measures of its respective large trading partner. Moreover, nominal linkages in two dimensional systems CA and US, as well as UK and US increase if agents perceive FX rates as rather stable. Therefore, monetary autonomy might vary according to public beliefs concerning the relevance of exchange rate targets in policy functions of the BoC or BoE (Lubik and Schorfheide 2007).

The remainder of the paper is organized as follows: The next Section introduces the data and sketches the liquidity adjustment of breakeven inflation. Section 3 provides a VAR analysis for international breakeven rates. As we cannot identify a viable long run relationship, monthly realized correlations of breakeven rate changes are determined in Section 4. Moreover, these correlation paths are conditioned on measurable economic variables (factors) characterizing *i*) global (real) shocks, *ii*) dominance of foreign monetary policy and *iii*) FX rate flexibility. Section 5 summarizes the main findings and concludes.

## 5.2 Liquidity Adjusted Breakeven Rates

### 5.2.1 Data

We analyze breakeven inflation rates as implied by the spread of 10 year nominal and inflation indexed government bond yields for a subsystem of the G7 comprising CA, FR, the UK and the US. Daily price quotes cover the time period 4/2/2001-9/30/2008. For the construction of 10 year breakeven inflation rates, 10 year nominal interest rates are available for all markets considered. In contrast, official quotes for 10 year real interest rates are only available for the UK and, since January 2003 published by the Federal Reserve Bank of St. Louis for the US. UK real rates as quoted by the BoE are directly drawn from Datastream. For the period before 2003, the US 10 year constant maturity rate is determined by linear interpolation of real yields to maturity

of 2011 and 2012 US TIPS yields. Accordingly, the 10 year constant maturity French real interest rate is approximated by linear interpolation of real yields to maturity of 2010, 2015 and 2030 OATis as obtained from the website of Agence France Trésor. The Canadian real rate is approximated by the long term benchmark bond yield reported by the BoC. Further details on the data sources are provided in Table 5.1.

## 5.2.2 Liquidity adjustment

Recent studies (Shen and Corning 2001, Sack and Elsasser 2004, or Shen 2006) suggest that 'raw' breakeven inflation rates might only weakly approximate expected inflation since premia paid for relatively lower liquidity shift the real yields of TIPS above the 'true' real interest rates. As a consequence, breakeven inflation rates tend to underestimate expected inflation and inflation risk. Owing to limited data availability we correct breakeven rates for liquidity risk on a quarterly basis. In the following, we use indices  $\tau$ ,  $t$  and  $m$  to refer to the quarterly, monthly and daily frequency of time series variables, respectively. Along the lines in Shen (2006), consider the regression

$$be_{\tau}^{(i)} - \pi_{\tau}^{e(i)} = \phi_1^{(i)} + \phi_2^{(i)} Lique_{\tau}^{(i)} + e_{\tau}^{(i)}, \quad (5.1)$$

where  $be_{\tau}^{(i)}$  and  $\pi_{\tau}^{e(i)}$  denote the breakeven inflation and the expected inflation rate in country  $i$  and quarter  $\tau$  with respect to a ten year horizon. Due to limited data availability, ten year expected inflation rates are approximated by the 10 year (US), 5 year (FR) and 1 year (CA, UK) survey based forecasts. Moreover,  $Lique_{\tau}^{(i)}$  is a TIPS market specific measure of the long run trend in market turnover obtained by Hodrick-Prescott (HP) filtering (Hodrick and Prescott 1997) in case of CA and the UK. Since corresponding turnover measures are not available for FR and the US, the HP implied long run change in volume of outstanding TIPS is used to approximate the evolution of liquidity of these markets. If liquidity measures were available at the monthly (daily) frequency, one could use (5.1) to adjust breakeven rates at these frequencies. Presuming that liquidity premia adjust only smoothly over time, monthly observations ( $Lique_t^{(i)}$ ) are determined from quarterly quotes ( $Lique_{\tau}^{(i)}$ ) by means of a frequency conversion implemented in EViews 5.1 ('quadratic match average'). Given estimates  $\hat{\phi}_1^{(i)}, \hat{\phi}_2^{(i)}$  from the (quarterly) regression (5.1), the monthly breakeven rate adjusted for liquidity is approximated as  $\tilde{be}_t^{(i)} = be_t^{(i)} - \hat{\phi}_1^{(i)} - \hat{\phi}_2^{(i)} Lique_t^{(i)}$ . Note that breakeven rates are only adjusted by means of a constant and a smooth liquidity measure. Therefore, the inconsistent time horizons of survey inflation data should not severely derogate the reliability of adjusted breakeven rates as long as short and

Table 5.1: Data sources

Data	Source (Code)
CA real l-term	Datastream (CNBBRLT)
CA 10y nom	<a href="http://www.bank-banque-canada.ca/en/rates/yield_curve.html">http://www.bank-banque-canada.ca/en/rates/yield_curve.html</a>
FR OATI 2009, 2013, 2030	<a href="http://www.aft.gouv.fr/article_1004.html?id_article=1004&amp;id_rubrique=256">http://www.aft.gouv.fr/article_1004.html?id_article=1004&amp;id_rubrique=256</a>
FR 10y nom	Datastream ( FRBRYLD)
UK 10y real	Datastream (UKGR10Y)
UK 10y nom	Datastream (UKBN10Y)
US TIPS 2011, 2012	<a href="http://research.stlouisfed.org/fred2/categories/82?cid=82&amp;pageID=2">http://research.stlouisfed.org/fred2/categories/82?cid=82&amp;pageID=2</a>
US 10y real	<a href="http://research.stlouisfed.org/fred2/series/DFII10?cid=82">http://research.stlouisfed.org/fred2/series/DFII10?cid=82</a>
US 10y nom	<a href="http://research.stlouisfed.org/fred2/series/DGS10?cid=115">http://research.stlouisfed.org/fred2/series/DGS10?cid=115</a>
CA WES 1y exp. infl.	Datastream (CNIFINFRR)
ECB SPF 5y exp. infl.; Disp.	<a href="http://www.ecb.int/stats/prices/indic/forecast/html/index.en.html">http://www.ecb.int/stats/prices/indic/forecast/html/index.en.html</a>
UK WES 1y exp. infl.	Datastream (UKIFINFRR)
US LSPF 10y exp. infl.; Disp.	<a href="http://www.philadelphiafed.org/research-and-data/real-time-center/livingston-survey/">http://www.philadelphiafed.org/research-and-data/real-time-center/livingston-survey/</a>
CA RRB av. market trading	Bank of Canada, Banking and Financial Statistics
FR OATI outstanding	<a href="http://www.aft.gouv.fr/article_1004.html?id_article=1004&amp;id_rubrique=256">http://www.aft.gouv.fr/article_1004.html?id_article=1004&amp;id_rubrique=256</a>
UK gilt market turnover	<a href="http://www.dmo.gov.uk/rpt_parameters.aspx?rptCode=D4J.2&amp;page=Turnover/Maturity">http://www.dmo.gov.uk/rpt_parameters.aspx?rptCode=D4J.2&amp;page=Turnover/Maturity</a>
US TIPS outstanding	Datastream(USGDMTSIA)
Crude oil WTI spot	Datastream (CRUDOIL)
Dow jones industrials	Datastream (DJINDUS)
Moodys AAA; BAA	Datastream (FRMCAAA; FRMCBAA)
EURO repo benchmark rate	Datastream(BBERB3M)
US federal funds rate	Datastream(FRFEDFD)
FX reserves	Datastream(CAI.1D.SA; EMI.1D.DA; UKI.1D.SA)
FX rates	Datastream(BBCADSP; UKECBSP, BBEURSP; USBRITP)



long term expected inflation rates roughly coincide on average.<sup>2</sup> In accordance with the Friedman-Ball hypothesis (Ball 1992) linking inflation uncertainty to inflation levels, inflation risk premia might fail time invariance implicitly assumed in (5.1). Thus, adjusted breakeven rates should slightly understate (overstate) expectations if these are low (high). Therefore, we regard breakeven rates as measures of 'future inflation prospects' or the 'inflation climate' rather than treating them as inflation expectations in a strict sense.

Liquidity adjusted breakeven inflation rates are displayed in Figure 5.1 (solid lines, upper four panels). As a benchmark, Figure 5.1 also shows the respective survey expected inflation rates (dashed lines). By graphical inspection adjusted breakeven rate dynamics appear consistent with survey implied inflation forecasts. In contrast to professional forecasters' long term inflation expectations in FR and the US, shorter term expected inflation rates in CA and the UK are stronger reflected in corresponding breakeven inflation rates. It appears that the market participants (as a whole) adjust long term expectations stronger to the current situation as it is the case for the subset of professional forecasters.<sup>3</sup>

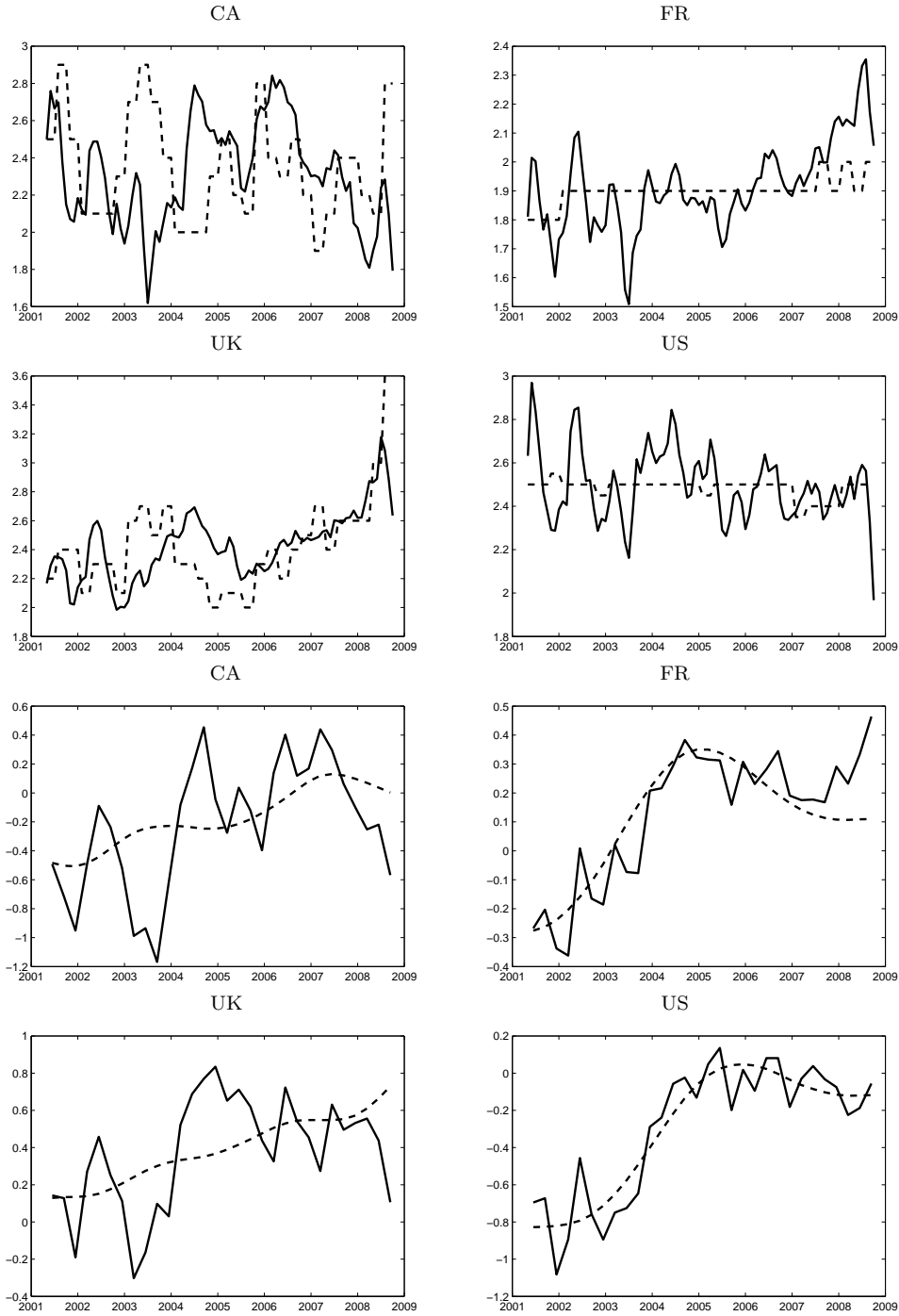
To give an impression of market liquidity,  $Liqu_{\tau}^{(i)}$ , and the differential between 'raw' breakeven rates and the corresponding survey data,  $be_{\tau}^{(i)} - \pi_{\tau}^{e(i)}$ , both are jointly sketched in the lower four panels of Figure 5.1. Since liquidity premia tend to push breakeven rates below true market expectations of inflation, we observe that breakeven rates have been increasing with the recent improvements in TIPS market liquidity. It is also noteworthy that most spreads  $be_{\tau}^{(i)} - \pi_{\tau}^{e(i)}$  are negative hinting at liquidity risk premia to be more important than inflation risk premia. Liquidity risk especially seems to matter for US markets since the estimated risk correction amounts up to 0.8 percentage points in absolute value. The left hand side panel of Table 5.2.2 provides static regression results and an ADF statistic for testing cointegration among  $be_{\tau}^{(i)} - \pi_{\tau}^{e(i)}$  and  $Liqu_{\tau}^{(i)}$ . An equilibrium relationship is diagnosed for CA, FR and the US, but not for the UK. This result might reflect that the UK indexed gilt market has been well developed throughout the entire sample period. However, we conduct liquidity adjustments in all cases.

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<sup>2</sup>Similar to liquidity risk adjustments one might also argue in favor of an adjustment of breakeven rates for inflation risk. Inflation risk premia, however, are not explicitly modeled since including survey implied cross sectional standard deviations of inflation in (5.1), the coefficient estimates suggested throughout negative or insignificant premia of inflation uncertainty. Consequently, we regard time stable inflation risk premia to be contained in the constant term in (5.1).

<sup>3</sup>Notably, also survey based evidence on long term expected inflation is not unique since e.g. the University of Michigan survey documents US long term expected inflation rates varying between 2.8 and 4.0 percentage points for the same sample period.

Figure 5.1: Breakeven inflation rates, expected inflation and liquidity risk premia



The upper four panels show monthly liquidity adjusted country specific breakeven inflation rates  $\tilde{be}_t^{(i)}$  (solid) versus the corresponding expected inflation from survey data (dashed line). The lower four panels show the quarterly difference  $be_\tau^{(i)} - \pi_\tau^{e(i)}$  (solid) versus the estimated risk premium  $\hat{\phi}_1^{(i)} + \hat{\phi}_2^{(i)} Liqu_\tau^{(i)}$  (dashed).

Table 5.2: Liquidity adjustment, integration and cointegration

	$\hat{\phi}_1$	$\hat{\phi}_2 \times E-03$	CI	ADF	$c$	Lags	$H_0$	5% CV	Trace
CA	-0.876 (-4.07)	0.190 (3.20)	-3.47	-2.902 (.126)	0.319 (2.81)	1	$CR = 0$	57.00	43.42 (.384)
FR	-0.578 (-8.21)	0.022 (10.07)	-3.25	-2.560 (.284)	0.265 (2.60)	2	$CR = 1$	29.50	20.49 (.439)
UK	-0.047 (-0.38)	0.877 (3.60)	-2.60	-2.961 (.097)	0.268 (2.96)	1	$CR = 2$	13.43	7.15 (.546)
US	-0.981 (-16.49)	0.052 (12.67)	-5.13	-5.365 (.000)	0.914 (5.30)	1	$CR = 3$	4.47	1.86 (.352)

The left hand side panel provides liquidity adjustment parameter estimates for the model  $be_\tau^{(i)} - \pi_\tau^{e(i)} = \phi_1 + \phi_2 Liqu_\tau^{(i)} + e_\tau^{(i)}$  ( $t$ -statistics in parentheses) based on quarterly data. Moreover, ADF statistics ( $p$ -values) for cointegration among  $be_\tau^{(i)} - \pi_\tau^{e(i)}$  and  $Liqu_\tau^{(i)}$  are given (third column). The corresponding 10% critical value of the Dickey-Fuller distribution is  $-3.07$ . The center panel shows ADF statistics for liquidity adjusted monthly breakeven rates  $\tilde{be}_t^{(i)}$ . Heteroskedasticity robust  $p$ -values in parentheses are obtained by resampling (Cavaliere and Taylor 2008). The right hand side panel documents trace test statistics (Johansen 1991) for the system of liquidity adjusted breakeven inflation rates and respective heteroskedasticity robust  $p$ -values in parentheses (Cavaliere, Rahbek and Taylor 2008).  $CR$  is the cointegration rank subject to testing. An intercept is included in the cointegrating relation and the VAR. Selection of lag order 2 accords with the AIC.

For further processing, the vector of monthly liquidity adjusted breakeven rates is denoted as  $y_t = (\tilde{be}_t^{(CA)}, \tilde{be}_t^{(FR)}, \tilde{be}_t^{(UK)}, \tilde{be}_t^{(US)})'$ . For convenience, we subsequently use the term 'breakeven inflation' instead of 'liquidity adjusted breakeven inflation'.

### 5.2.3 Economic significance

All breakeven rates displayed in Figure 5.1 show marked time variation and even trending patterns (FR or UK). Given credible monetary policy strategies, one would rather expect four stationary processes with mean close to the respective long run inflation target. For the case of the UK, a trending breakeven rate might be realistic since the BoE changed its inflation target in 2003. The BoE, however, reduced the target from 2.5 to 2.0 percentage points while Figure 5.1 indicates increasing breakeven inflation. The BoC's inflation target range is 1 to 3 per cent, while it aims to keep inflation at 2 per cent. Although the target range is met throughout, one observes substantial deviations from the 2% target according to survey and breakeven inflation implied expectations. Similarly, also the evolution of the US breakeven inflation rate hints at perceptions about future inflation and associated risks which are rather unstable in

the short run. In the medium term, however, US inflation expectations appear to be well anchored. Underpinning the high reputation of the ECB, the French breakeven inflation rate varies with small amplitude and indicates rather stable expectations.

## **5.3 First Order Analysis of Breakeven Inflation**

In this section we describe dynamic and unconditional features of breakeven inflation rates. In the first place, unit root tests are applied to investigate if breakeven rates are stationary around a certain anchor. Secondly, we address conditional mean dynamics by means of a VAR model. Estimation results and diagnostic statistics for conditional mean modeling are documented in Table 5.2.2 and Table 5.3. If not stated explicitly, the discussion of empirical results refers to the 5% significance level.

### **5.3.1 (Co)integration tests**

As a preliminary analysis we conduct ADF tests for the four breakeven rates applying heteroskedasticity robust critical values obtained by resampling (Cavaliere and Taylor 2008). Unit root test statistics in the center panel of Table 5.2.2 indicate nonstationarity of CA, FR and UK breakeven inflation rates while the US rate is diagnosed stationary. Assuming nonstationarity of breakeven rates, trace statistics for testing cointegration and corresponding heteroskedasticity robust critical values (Cavaliere, Rahbek and Taylor 2008) are documented in the right hand side panel of Table 5.2.2. The trace statistic does not hint at cointegration such that the diagnosed system rank implies the prevalence of four independent stochastic trends. Thus, there is no evidence that expected inflation rates share a stable equilibrium relation. Given the current floating regimes, the smaller economies (CA and the UK) should be able to autonomously determine the (expected) inflation rate targets in the long run. Full monetary independence, however, is not implied by stochastic trending since changes of long term expected inflation rates and risk premia could be correlated across markets in the short and medium term.

### **5.3.2 VAR estimates and Granger causality tests**

We address the issue of nominal transmission among international breakeven rates by means of a VAR in first differences, augmented with some control variables to discriminate between common reactions to global (real) shocks and nominal transmission. To measure common dynamics in response to (real) global shocks, we employ lagged log

changes of the West Texas Intermediate (WTI) oil price index and log returns of the Dow Jones Industrial Average.<sup>4</sup>

VAR parameter estimates are documented in Table 5.3. Heteroskedasticity consistent  $t$ -ratios indicate that lagged changes in the US inflation climate affect breakeven rates in CA and (with 10% significance) in FR. In turn, the Northern American inflation climate seems to be positively affected by the first lag of EMU breakeven rate changes. Cross country transmission of changes in future inflation prospects appears moderate, since the corresponding degrees of partial explanation for the VAR equations range from 6.4% (UK) to 22.8% (CA).

To underpin the significance of these transmission channels, heteroskedasticity consistent Granger causality test statistics (Hafner and Herwartz 2009) are also shown in Table 5.3. These confirm rates in CA to be Granger caused by FR, UK and US rates, while French rates are affected by the UK and (with 10% significance) by the US inflation climate. Interestingly, also the US inflation climate is Granger caused by changes of breakeven inflation in CA. Thus, a unique pattern of Granger causality reflecting nominal US or EMU dominance over CA and the UK cannot be diagnosed. In a similar vein, Yang et al. (2006) find that inflation is transmitted not only from large economies to smaller ones, but also in the opposite direction.

Our findings suggest that sentiments about future inflation rates do not evolve independently in an international context despite of free floating FX rates. At least in the medium run, inflation expectations interact, although monetary policy actions are conducted with respect to domestic goals. This could indicate that for small open economies, domestic and foreign monetary policies influence long term inflation expectations in the short and medium run. Under these circumstances central banks could temporarily face difficulties to maintain fixed domestic long run ex-ante inflation rates. Moreover, the existence of multidirectional transmission might highlight a need of strengthened coordination of monetary policies.

## 5.4 Second Order Analysis of Breakeven Inflation

While first order analysis provides evidence for absence of long term linkages, full monetary independence, however, might not prevail since breakeven inflation rate changes might be correlated across markets. In this section, the case of short run monetary independence is investigated focusing at second order dynamics. Firstly,

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<sup>4</sup>Initially, we also considered changes in the US corporate bond spread as implied by the difference in yields of Moody's BAA and AAA bond indices. Since this measure did not significantly impact on breakeven inflation dynamics, it is not included in the VAR.

Table 5.3: VAR estimate and Granger causality

	$\Delta\pi^{e(CA)}$	$\Delta\pi^{e(FR)}$	$\Delta\pi^{e(UK)}$	$\Delta\pi^{e(US)}$		$\Delta\pi^{e(CA)}$	$\Delta\pi^{e(FR)}$	$\Delta\pi^{e(UK)}$	$\Delta\pi^{e(US)}$
CA(-1)	0.086 (0.68)	-0.071 (-1.03)	-0.086 (-1.10)	0.018 (0.18)	<i>DOW</i> (-1)	0.116 (0.31)	0.845 (4.81)	0.973 (2.99)	0.743 (2.17)
CA(-2)	-0.122 (-0.91)	-0.048 (-0.76)	-0.018 (-0.19)	-0.252 (-2.83)	c	-0.009 (-0.71)	-0.002 (-0.32)	-0.002 (-0.19)	-0.014 (-1.37)
FR(-1)	0.674 (2.41)	0.130 (1.09)	0.280 (1.21)	0.516 (2.10)	$R^2$	.363	.508	.329	.412
FR(-2)	-0.200 (-0.76)	-0.192 (-1.34)	0.160 (1.04)	-0.102 (-0.45)	$R^2_{part}$	.228	.180	.064	.214
UK(-1)	0.072 (0.34)	0.129 (1.38)	0.252 (1.53)	0.142 (0.93)		$\Delta\pi^{e(CA)}$	$\Delta\pi^{e(FR)}$	$\Delta\pi^{e(UK)}$	$\Delta\pi^{e(US)}$
UK(-2)	-0.462 (-2.66)	-0.292 (-2.76)	-0.192 (-1.26)	-0.112 (-0.66)	CA $\leftrightarrow$		2.07 (.355)	1.25 (.535)	8.05 (.018)
US(-1)	0.012 (0.07)	0.102 (1.20)	0.027 (0.22)	0.029 (0.22)	FR $\leftrightarrow$	6.36 (.042)		2.21 (.332)	4.54 (.103)
US(-2)	0.424 (2.56)	0.172 (1.65)	-0.028 (-0.20)	-0.054 (-0.36)	UK $\leftrightarrow$	7.13 (.028)	9.97 (.007)		1.43 (.490)
<i>OIL</i> (-1)	0.239 (1.18)	0.325 (3.35)	0.158 (1.46)	0.190 (1.48)	US $\leftrightarrow$	6.63 (.036)	5.47 (.065)	0.01 (.953)	
					All $\leftrightarrow$	33.37 (.000)	17.37 (.008)	4.69 (.585)	21.03 (.002)

VAR parameter estimates (left and upper right panel) as well as Granger causality tests (lower right) for a system of liquidity adjusted breakeven rates in first differences. Dependent (explanatory) variables are ordered by column (row). Lag selection is according to the AIC. Residuals of the model are serially uncorrelated according to a respective LM-statistic: 12.76,  $p$ -value 0.690 (Breusch 1978, Godfrey 1978). VAR  $t$ -statistics, as well as  $p$ -values for tests on Granger non-causality in parentheses are based on heteroskedasticity consistent covariance estimates (White 1980, Hafner and Herwartz 2009).  $R^2_{part}$  is the degree of explanation for  $\Delta\pi_t^{e(i)}$ , after partialling out the effects of  $\Delta\pi_{t-1}^{e(i)}$ ,  $\Delta\pi_{t-2}^{e(i)}$ , *OIL*(-1) and *DOW*(-1).

the concept of monthly realized (co)variance estimation is briefly sketched. Then, empirical patterns of realized correlations of international breakeven inflation rates are discussed. Turning to the issue of state dependence, we list economic factors which potentially carry explanatory content for the joint evolution of realized monthly breakeven correlations. Lastly, potential state dependence of breakeven correlations is investigated by means of a seemingly unrelated regression (SUR) model (Zellner 1962).

### 5.4.1 Realized correlations

Arguably, one of the simplest ways to assess volatilities of diffusion processes is to aggregate over their high frequency (realized) squared increments (French, Schwert and Stambaugh 1987, Schwert 1989, Andersen, Bollerslev, Diebold and Labys 2003). In this nonparametric setting the consistent estimator of the monthly breakeven (co)variance is the sum over squared daily increments,<sup>5</sup> i.e.

$$RV_{ij,t} = \sum_{m \in t} \Delta \tilde{b}e_m^{(i)} \Delta \tilde{b}e_m^{(j)}, \quad \{i, j\} \in \{\text{CA, FR, UK, US}\}. \quad (5.2)$$

Monthly realized correlations are accordingly

$$\rho_t(i, j) = \frac{RV_{ij,t}}{\sqrt{RV_{ii,t}RV_{jj,t}}}, \quad i \neq j. \quad (5.3)$$

For the empirical analysis, we focus on the set of correlations  $\rho_t(\text{CA,US})$ ,  $\rho_t(\text{FR,UK})$ ,  $\rho_t(\text{FR,US})$  and  $\rho_t(\text{UK,US})$  and disregard  $\rho_t(\text{CA,UK})$  and  $\rho_t(\text{CA,FR})$ . Opposite to the latter, the former set represents linkages of integrated economies to large trading partners which might be dominating in nominal terms.

### 5.4.2 Empirical correlation patterns

Realized monthly correlations of breakeven rates are displayed in Figure 5.2. Moreover, corresponding unconditional means, standard deviations, test statistics for normality and differences in mean are given in Table 5.4.

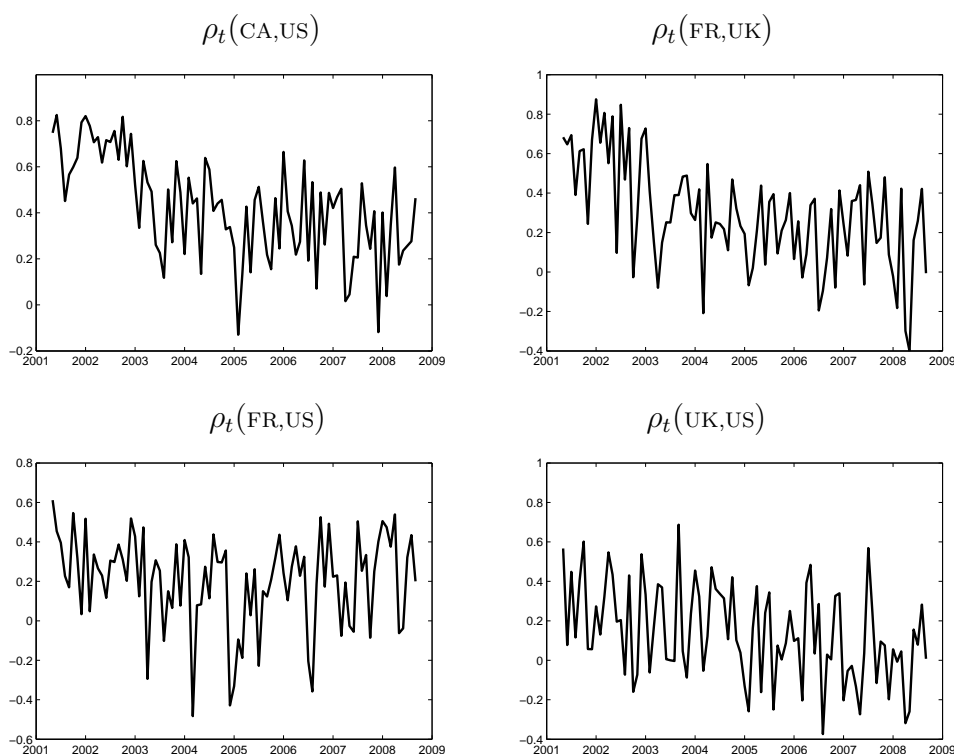
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<sup>5</sup>For the calculation of monthly realized correlations, daily changes of breakeven rates are used.

Daily breakeven rates are calculated as  $\tilde{b}e_m^{(i)} = b e_m^{(i)} - \hat{\phi}_1^{(i)} - \hat{\phi}_2^{(i)} \text{Liqu}_m^{(i)}$ , where the daily liquidity measure is obtained by means of the frequency conversion mentioned before. The high frequency increments used to calculate monthly realized correlations of breakeven inflation rates are based on daily changes, i.e.  $\Delta \tilde{b}e_m^{(i)} = \tilde{b}e_m^{(i)} - \tilde{b}e_{m-1}^{(i)}$ .

Eyeball inspection of Figure 5.2 reveals low persistence as well as a high amplitude of breakeven correlations. For example,  $\rho_t(\text{CA,US})$  varies between 0.83 and  $-0.13$ , its sample average is  $\bar{\rho}(\text{CA,US}) = 0.43$ , where  $\bar{\rho}(i, j) = \frac{1}{T} \sum_{t=1}^T \rho_t(i, j)$ . The Northern American subsystem features the highest correlations, followed by the European subsystem.

Figure 5.2: Monthly realized breakeven correlations



Time paths of the monthly realized correlation matrix elements  $\rho_t(i, j)$  defined in equation (5.3).

Despite being characterized by a mean reverting tendency in the short run, correlations seem to change in the medium run in terms of their average levels. For example, correlations  $\rho_t(\text{CA,US})$ ,  $\rho_t(\text{FR,UK})$  and  $\rho_t(\text{UK,US})$  have been decreasing since 2001. In particular, rates in CA and the US show a decrease in correlations in the sequel of spring 2003 which coincides with strong depreciations of the US Dollar.

Mean difference test results given in Table 5.4 confirm  $\bar{\rho}(\text{CA,US})$  to be significantly higher than  $\bar{\rho}(\text{FR,UK})$ , while  $\bar{\rho}(\text{FR,UK})$  exceeds  $\bar{\rho}(\text{FR,US})$  and  $\bar{\rho}(\text{UK,US})$ . The inflation targeting policy of the BoC could be impaired since the domestic long run inflation



climate varies with expectations of long run US sentiments and, thus, does not fully reflect domestic targets. Similarly, FR and the UK are subject to common shocks hitting inflation prospects but, as empirical correlations are significantly lower, the BoE is likely better able to target domestic inflation than the BoC.

Table 5.4: Selected statistics of realized correlations and tests for equality

	Mean	Std. Dev.	JB	$\rho_t(\text{FR,UK})$	$\rho_t(\text{FR,US})$	$\rho_t(\text{UK,US})$
$\rho_t(\text{CA,US})$	.429	.225	1.74 (.419)	3.61 (.000)	6.38 (.000)	8.22 (.000)
$\rho_t(\text{FR,UK})$	.295	.266	0.30 (.862)	-	2.24 (.026)	3.94 (.000)
$\rho_t(\text{FR,US})$	.207	.234	12.10 (.002)	-	-	1.84 (.067)
$\rho_t(\text{UK,US})$	.145	.242	1.93 (.381)	-	-	-

Descriptive statistics for monthly realized correlations of breakeven inflation changes (left hand side panel), and conventional  $t$ -tests for equality in mean (right hand side), with  $p$ -values in parentheses. For instance, the last element in the first row documents the  $t$ -statistic for testing the hypothesis  $\rho(\text{CA,US}) = \rho(\text{UK,US})$ .

Ex-ante reasoning suggests that in case of free floating FX regimes, high shock interdependence of long term inflation prospects should arise in response to global shocks. However, as the actual evolvment of realized correlations appears distinct over groups ( $\rho_t(\text{CA,US})$  and  $\rho_t(\text{FR,UK})$  vs.  $\rho_t(\text{FR,US})$  and  $\rho_t(\text{UK,US})$ ), one might conjecture that the observed correlation patterns are also reflections of 'local' signalling from a large, the US say, to small economies, as CA or the UK.

In the long run, under free floating FX rates expected inflation should be completely determined by the expected growth in monetary aggregates. Observing correlation measures up to 0.8 questions either full monetary autonomy in presence of floating FX rates or the prevalence of a 'pure float' in the spirit of Calvo and Reinhart (2002). We subsequently introduce state variables to uncover potential triggers of nominal linkage.

### 5.4.3 State dependence of correlations

#### Factors used to address state dependence

Three types of conditioning variables are distinguished that could generate interdependence of long term inflation prospects. In the first place, international linkage could reflect the dominance of global over domestic shocks (Longin and Solnik 1995). To identify comovements caused by common exposure to global (real) risk factors, we employ measures such as oil price volatility, the US high yield corporate bond spread and US stock market volatility. Formally, these variables are defined as

$$\sigma_t^{oil} = \sqrt{RV_{t-1}^{oil}}, \quad RV_t^{oil} = \sum_{m \in t} (\Delta \ln OIL_m)^2, \quad (5.4)$$

$$cbs_t = \xi_{t-1}^{B,US} - \xi_{t-1}^{A,US}, \quad (5.5)$$

$$\sigma_t^{dow} = \sqrt{RV_{t-1}^{dow}}, \quad RV_t^{dow} = \sum_{m \in t} (\Delta \ln DOW_m)^2, \quad (5.6)$$

where  $OIL_m$  is the daily price for West Texas Intermediate (WTI) crude oil,  $\xi_t^{A,US}$  ( $\xi_t^{B,US}$ ) denotes Moodys aggregate AAA (BAA) US corporate bond yields, and  $DOW_m$  is the daily 'Dow Jones Industrials Average' index. Secondly, nominal shock interdependence of long run inflation prospects might reflect that large economies dominate their trading partners. Corresponding to core policy instruments, realized volatilities of the three month *Euro Repo Benchmark Rate* and the *Federal Funds Rate* are seen to potentially trigger dependence of expected inflation in this respect. Reflecting the current state of monetary policy in the EMU (US) the spread between FR (US) ten year nominal interest rates and the policy rates mentioned before are further conditioning variables. Moreover, we consider EMU (US) inflation uncertainties approximated by the cross sectional dispersion of respective long term inflation surveys. Formally, the second set of variables is

$$\sigma_t^{sti,k} = \sqrt{RV_{t-1}^{sti,k}}, \quad RV_t^{sti,k} = \sum_{m \in t} (\Delta sti_m^k)^2, \quad (5.7)$$

$$spr_t^k = lti_{t-1}^k - sti_{t-1}^k, \quad (5.8)$$

$$\sigma_{t \in \tau}^{disp,k} = \sqrt{CSD_{\tau-1}^k}, \quad CSD_{\tau}^k = \frac{1}{S_{\tau}^k - 1} \sum_{s=1}^{S_{\tau}^k} (\pi_{\tau}^{e(k,s)} - \pi_{\tau}^{e(k)})^2, \quad (5.9)$$

where  $sti_t^k$  ( $lti_t^k$ ) denotes the monthly short term (long term) interest rate in country  $k$ ,  $k \in \{\text{US, EMU}\}$ . Similarly,  $sti_m^k$  refers to daily measures. Moreover,  $\pi_{\tau}^{e(k,s)}$  denotes an individual survey forecast of participant  $s$  in country  $k$  and quarter  $\tau$ , and  $S_{\tau}^k$  is the

number of professional forecasters in quarter  $\tau$ ,  $k \in \{\text{US}, \text{EMU}\}$ .<sup>6</sup> The third group of variables accounts for the view that small economies face nominal dependencies since they refrain from pursuing a 'pure' float (Calvo and Reinhart 2002). We rely on the fluctuations in FX reserves and the realized volatility of FX rates to characterize the state of floating. Floating state measures are defined as

$$\sigma_t^{res,k} = \sqrt{RV_{t-1}^{res,k}}, \quad RV_t^{res,k} = \sum_{l=1}^3 (\Delta RES_{t-l}^k)^2, \quad (5.10)$$

$$\sigma_t^{fx,ij} = \sqrt{RV_{t-1}^{fx,ij}}, \quad RV_t^{fx,ij} = \sum_{m \in t} (\Delta \ln FX_m^{ij})^2, \quad (5.11)$$

where  $RES_t^k$  denotes monthly FX reserves of country  $k$ ,  $k \in \{\text{CA}, \text{UK}, \text{EMU}\}$ . Moreover,  $FX_m^{ij}$  are daily units of floating currency  $i$  per one unit of currency  $j$ ,  $\{i, j\} \in \{\text{CA}, \text{US}; \text{UK}, \text{FR}; \text{FR}, \text{US}; \text{UK}, \text{US}\}$ .

Now, consider the regression

$$\begin{aligned} \rho_t(i, j) = & \beta_1 + \beta_2 \sigma_t^{oil} + \beta_3 cbs_t + \beta_4 \sigma_t^{dow} + \beta_5 \sigma_t^{sti,us} + \beta_6 \sigma_t^{sti,emu} + \beta_7 spr_t^{us} + \\ & \beta_8 spr_t^{emu} + \beta_9 \sigma_t^{disp,us} + \beta_{10} \sigma_t^{disp,emu} + \beta_{11} \sigma_t^{res,i} + \beta_{12} \sigma_t^{fx,ij} + e_t^{(ij)}, \end{aligned} \quad (5.12)$$

with  $\{i, j\} \in \{\text{CA}, \text{US}; \text{UK}, \text{FR}; \text{FR}, \text{US}; \text{UK}, \text{US}\}$ . Throughout referring to information available in month  $t - 1$ , explanatory variables in (5.12) are predetermined such that the regression model may also serve as an ex-ante forecasting design. Given marked patterns of cross equation correlation of  $e_t^{(ij)}$ , the model in (5.12) is implemented as a SUR system. Noting that this system comprises a total of 48 parameters, coefficients are subject to sequential pretesting. Excluding the intercept terms those parameter estimates are subsequently removed from the system that obtain the smallest  $t$ -ratio in absolute value. The test sequence terminates once all parameter estimates share  $t$ -statistics that are at least unity in absolute value to keep the likelihood of joint insignificance of removed parameters at conventional nominal levels.

### Estimation results

Parameter estimates for (5.12) are given in Table 5.5. Residual diagnostics indicate absence of first order serial correlation, homoskedasticity and, for the case of modeling  $\rho_t(\text{CA}, \text{US})$  and  $\rho_t(\text{UK}, \text{US})$ , unconditional normality. Accordingly, parameter estimates

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<sup>6</sup>When calculating cross sectional standard deviations, we discarded an 'outlier' in the US survey data in 2005Q3, where one participant posted a 6% expected inflation while the remaining 36 participants declared expectations within a range of 2% - 4%.

are likely efficient and  $p$ -values reliable.

Concerning the impact of global (real) shocks on correlation dynamics, we observe the volatility of the US stock market to raise realized correlations  $\rho_t(\text{CA,US})$  and  $\rho_t(\text{FR,US})$ . Moreover,  $\rho_t(\text{UK,US})$  decreases in response to an increase in oil price volatility. However, the overall impact of global (real) shock variables ( $cbs_t$ ,  $\sigma_t^{oil}$  and  $\sigma_t^{dow}$ ) is moderate.

Turning to the explanatory content of variables summarizing monetary policy ( $\sigma_t^{sti,k}$ ,  $spr_t^k$  and  $\sigma_{t \in \tau}^{disp,k}$ ), we observe that perceptions of future inflation are more synchronous in CA and the US if the US monetary policy tries to stimulate the economy by means of setting low policy rates inducing a steep yield curve. In addition, high inflation uncertainty in large economies raises the respective correlations with its smaller trading partners. Inflation uncertainty in the US impacts positively on correlations  $\rho_t(\text{CA,US})$  and  $\rho_t(\text{UK,US})$ . Similarly, EMU inflation uncertainty fosters  $\rho_t(\text{FR,UK})$ . Weakened credibility of central banks or insufficiently anchored inflation expectations in major economies seem to matter for the trading partners' inflation climate. The latter fact might underline that a certain share of observed correlation dynamics reflects comovements of inflation risk premia, which also contribute to nominal dependence.

Impact estimates of measures of perceived floating ( $\sigma_t^{res,k}$  and  $\sigma_t^{fx,ij}$ ) suggest that realized correlations  $\rho_t(\text{UK,US})$  are increasing in the volatility of the UK FX reserve positions with 10% significance. Lastly, a significantly negative relationship is found between  $\rho_t(\text{CA,US})$  and respective FX volatilities. The impact estimate for FX rate flexibility in the European subsystem is also negative, but insignificant. Thus, nominal transmission effects operating between the US on the one hand and CA and the UK on the other hand respond to perceived excess FX rate stability. Agents noting an increased stability of FX rates or activity of the domestic central bank on FX markets appear to adapt their behavior such that linkages among breakeven inflation increase. To some extent, time variation of breakeven correlations can be traced back to changing beliefs with regard to the actual float. The agents' uncertainty, in turn, reflects the fact that central banks hosting a floating currency might discretionary intervene in FX markets.

Economically, CA has a strong incentive to influence the FX rate shared with the US in light of its trade relations. The dominance of the US in the Canadian trade portfolio can be verified by noting that the US Dollar enters with approximately 75% the Canadian trade weighted exchange rate index. Moreover, the BoC is known to eventually intervene in FX markets in both coordinated and non coordinated fashion.

Table 5.5: Conditional analysis of monthly realized correlation dynamics-SUR

Var	$\rho_t(\text{CA,US})$	$\rho_t(\text{FR,UK})$	$\rho_t(\text{FR,US})$	$\rho_t(\text{UK,US})$	Var	$\rho_t(\text{CA,US})$	$\rho_t(\text{FR,UK})$	$\rho_t(\text{FR,US})$	$\rho_t(\text{UK,US})$
c	0.197 (1.16)	-0.515 (-1.86)	0.182 (2.50)	-0.539 (-1.99)	$spr_t^{us}$	0.051 (3.15)			0.032 (1.09)
$cbs_t$		0.168 (1.30)			$spr_t^{emu}$		0.068 (1.21)	-0.062 (-1.83)	0.076 (1.11)
$\sigma_t^{oil}$	-0.931 (-1.51)			-1.569 (-2.03)	$\sigma_t^{\pi^e,us}$	0.752 (2.53)	0.925 (1.77)		1.259 (2.55)
$\sigma_t^{dow}$	3.302 (3.38)	2.238 (1.45)	2.568 (2.24)	1.575 (1.15)	$\sigma_t^{\pi^e,emu}$	0.659 (1.86)	1.386 (2.74)		0.688 (1.52)
$\sigma_t^{sti,us}$	0.143 (1.40)	0.180 (1.27)			$\sigma_t^{res,i}$		-3.4E-05 (-1.31)		4.4E-05 (1.80)
$\sigma_t^{sti,emu}$	-0.138 (-1.00)	-0.338 (-1.63)	-0.224 (-1.19)	-0.277 (-1.49)	$\sigma_t^{fx,j}$	-32.44 (-5.78)	-12.53 (-1.14)		
DW	2.133	1.815	1.760	1.850	Wh	.110	.131	.703	.848
JB	.733	.006	.003	.767					

Parameter estimates for the SUR system regression  $\rho_t(i, j) = \beta_1 + \beta_2 cbs_t + \beta_3 \sigma_t^{oil} + \beta_4 \sigma_t^{dow} + \beta_5 \sigma_t^{sti,us} + \beta_6 \sigma_t^{sti,emu} + \beta_7 spr_t^{us} + \beta_8 spr_t^{emu} + \beta_9 \sigma_t^{\pi^e,us} + \beta_{10} \sigma_t^{\pi^e,emu} + \beta_{11} \sigma_t^{res,i} + \beta_{12} \sigma_t^{fx,ij} + e_t^{(ij)}$ , with  $\{i, j\} \in \{\text{CA,US}; \text{UK,FR}; \text{FR,US}; \text{UK,US}\}$ . Variables with the smallest  $t$ -statistics are successively removed until all  $t$ -statistics exceed unity in absolute value. Residual diagnostics on first order serial correlation (Durbin Watson statistic, DW), results for White's heteroskedasticity test (Wh, White 1980) and the Jarque-Bera test (JB) on unconditional normality are given for each equation in the lower panel ( $p$ -values).

Occasional interventions are discretionary, rather than systematic.<sup>7</sup> In addition to a rather opaque commitment to exchange rate flexibility, such apparent incentives might be one reason for the sensitivity of CA long term nominal linkage to observed exchange rate stability.

In contrast, it appears more difficult to assess if the UK inflation climate is more subject to US or EMU monetary policy since nominal dependence seems to be governed by the exposure to both EMU and US inflation uncertainties and UK/US FX rate stability. However, as average correlations  $\bar{\rho}_{(FR,UK)}$  and  $\bar{\rho}_{(UK,US)}$  are significantly smaller than  $\bar{\rho}_{(CA,US)}$ , future inflation prospects in the UK are less subject to US or EMU monetary policies in comparison with the case of CA.

In summary, our results illustrate that integrated small economies with de jure flexible FX rate regimes might occasionally lack full control over domestic long term inflation expectations. Free floating currencies per se do not guarantee complete control over long term inflation expectations, especially in states of perceived FX rate stability and inflation uncertainty prevailing in major economies. Such state characteristics of nominal linkages should be taken into account by central banks conducting inflation targeting policy.

## 5.5 Conclusions

By standard macroeconomic theory floating FX rates establish that monetary authorities have full control over the domestic price level or the target inflation rate. Questioning full autonomy, we study first and second order linkages of liquidity adjusted breakeven inflation rates interpreted as a measure of future inflation prospects in CA, FR (the EMU), the UK and the US. Apart from first order causality testing, the dynamics of realized comovements are related to measurable economic states accounting for the monetary policy impact of a foreign (large) economy on domestic inflation prospects.

Liquidity adjusted breakeven inflation rates are mostly nonstationary, but not subject to a long run relation. The first order causality structure indicates that future inflation prospects are transmitted not only from large economies (such as the US) to smaller economies, but also in the reverse direction. Second order dynamics uncover that inflation prospects are characterized by markedly time varying comovements. Generally, stronger contemporaneous correlations are observed for integrated

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<sup>7</sup>See the homepage of the BoC (<http://www.bankofcanada.ca/en/backgrounders/bg-e2.html>) for further information

economies sharing intense trade relationships. Highest correlation statistics are found for CA and US breakeven inflation rate changes, while also correlations between the UK and the EMU are considerably high.

Finding that relatively small integrated economies such as CA show nominal linkage to their major trading partner, we try to discover fundamental determinants of nominal interdependence. Comovements are related to (observable) characteristics of the current conduct of monetary policy such as the volatilities in policy short term interest rates, yield curve slopes and inflation uncertainty prevailing in the US and the EMU. To motivate why the monetary policy in large economies could influence future inflation prospects in smaller economies, we further relate these linkages to the 'fear of floating' (Calvo and Reinhart 2002). Changes in breakeven correlations are seen to arise due to agents' uncertainty about the perfection of the current float. Inflation uncertainty in large economies is found to increase realized correlations shared with smaller trading partners. In addition, the empirical evidence supports a significantly negative relationship between realized correlations and realized FX rate volatilities among CA and the US. Moreover, an increase in the volatility of FX reserves held by the BoE leads to an uprise of UK and US breakeven correlations. Measures for global (real) shocks and financial turbulence such as high yield corporate bond spreads, oil price volatilities or stock market volatilities do moderately impact on observed correlation features.

In summary, although under floating FX rates long run inflation expectations should be determined solely by domestic monetary policy, international breakeven inflation rates feature considerable dependence in both first and second order moments. These findings have important monetary policy implications for small economies closely linked to large trading partners. Firstly, the monetary authorities in the former have to monitor inflation uncertainties in other dominant (global) economies to assess the impact on the domestic inflation climate. Secondly, a decrease in FX rate flexibility could imply that domestic inflation prospects are more sensitive to the foreign monetary policy. The evidence in this study provides additional guidance for the conduct of monetary policy, conditional on the relationship between FX rate volatility and inflation uncertainty on the one hand and comovements of breakeven inflation on the other hand. Under state dependence, the monetary authorities are naturally interested in a quantitative evaluation of first and second order transmission of expected inflationary pressures from abroad.

## Concluding Remarks

In this work it is argued that the most basic decisions of monetary policy can be motivated with regard to a few basic principles. For instance, the choice of a low inflation policy reflects the view that this helps to realize economic growth and stability in the medium to long run. The decision for capital accounts openness is motivated by presumed welfare gains from improved capital allocation. Lastly, exchange rate flexibility is seen as a prerequisite for being able to pursue a domestic monetary policy in presence of open capital markets.

Empirical studies on such assumptions, however, suffer from several conceptual shortcomings that might give rise to biased conclusions. Namely, studies which investigate the interrelationship between inflation, inflation uncertainty, output and output volatility might not provide representative evidence on typical interactions that might prevail in modern world economies. Studies that address capital allocation efficiency in terms of long run real interest rate convergence likely provide misleading results due to the use of biased measures for long run real interest rates. Lastly, empirical evidence on monetary autonomy suffers from overly stylized panel model specifications and potentially misleading quantifications of nominal dependence.

The aim of this empirical work is to improve on these aspects. To provide an estimate on the general interrelationship among inflation, output and its uncertainties, we aggregate information obtained from country specific VAR GARCH-M estimates of inflation and output dynamics in 34 economies during the current era of low inflation policies. Inferring from the cross section of parameter estimates in the framework of mean group estimation, we find that both inflation and inflation uncertainty tend to exert a negative influence on economic growth for the majority of considered economies. Hence, inflation and its uncertainty should be variables of concern for monetary authorities.

To test for capital allocation efficiency, we investigate the convergence properties of observable ex ante real interest rates obtained from Treasury Inflation Indexed Securities. Both inference based on cointegration and unit root tests suggests long run convergence of real capital market interest rates, and hence, the efficiency of



capital allocation for the considered subset of G7 countries.

Turning to the question of monetary independence, we employ a flexible semi-parametric panel methodology to express the tradeoff implied by the macroeconomic trilemma in a general manner. Evidence suggests that economies with fixed FX rates and closely integrated capital markets share one-to-one interest rate movements with their reference economies. In contrast, states that are either characterized by higher FX variation or lower capital mobility (or both) feature markedly less transmission.

Questioning the use of short run interest rates to measure monetary autonomy in general, I provide an alternative empirical assessment of monetary dependence under flexible exchange rates in terms of assessing comovements of international breakeven inflation rates. Evidence suggests that international breakeven inflation rates feature considerable dependence in both first and second order moments. Such comovements are empirically traced back to global states of inflation uncertainty and exchange rate (in)flexibility.

In summary, this work confirms the current view on the implications of monetary stability, open capital accounts and exchange rate flexibility.

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