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Real interest rate convergence among G7 countries*

Jana Riedel†

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Abstract

We analyze real interest rate convergence among six industrialized countries in between 1975M1-2011M3 within a multi-country framework by means of a dynamic latent factor model. The real interest rates are decomposed into permanent and transitory factors, and country-specific components. Time-varying variances allow for an endogenous transition from a high variance regime at the beginning of the sample towards a low variance regime. The estimation results suggest that four permanent and four transitory components capture the real interest rate dynamics among the sample of industrialized countries. The common component variances mostly decline over time, and in part even converge to values close to zero. This indicates a reduction in the number of stochastic components over time, which can be interpreted as confirmation of the convergence hypothesis. We observe rapid convergence during the late 1970s and 1980s, followed by slower transition since the mid-1990s when financial markets had already been highly integrated.

Keywords: latent dynamic factor model, unobserved components, real interest rate, RIP, financial market integration, G7

JEL classification: C32, F36, F32, E43

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1 Introduction

The behavior of real interest rates among countries has been and still is an important question in international monetary economics. Standard monetary exchange rate models (see Frenkel 1976; Dornbusch 1976; and Frankel 1979) are based on the assumption of real interest parity (RIP). This implies two conditions, perfect capital mobility and perfect price adjustment for goods, services and factors of production. Thus, if financial and goods markets are perfectly integrated, i.e., the uncovered interest rate parity (UIP) and the purchasing power parity (PPP) do hold jointly, real interest rates must equalize among countries.

Since the 1970s and early 1980s, several developed countries liberalized their financial markets. Capital controls were loosened in 1973 in the United States and Germany after the Bretton-Woods breakdown, later in 1976 in the United Kingdom (UK) and Japan, and in 1986 in Italy and France. Assuming risk neutrality, one would expect to observe a common world interest rate or at least real rate equalization tendencies, itself fostered by reduced information costs and the emergence of new financial instruments.

First empirical results point in a different direction. Standard regression-type test results imply that real rates have not been equalized across countries, at least until the early 1990s. In his seminal paper Mishkin (1984b) applies this approach to quarterly data from several OECD countries and rejects equality of real rates. Cumby & Mishkin (1986) analyze comovements in real rates among the US and several European countries. Real interest rates are increasing during the 1970s and 1980s. US rates are characterized by a higher level and increased variability after the monetary policy change in 1979 than before, and Canadian rates follow a similar pattern. Interestingly, in France and Germany the rise is much lower and real rates peaked with a delay in the late 1980s in Germany and even later in France.

Several authors (see, e.g., Mishkin 1984a; Cumby and Obstfeld 1984) use realized inflation data assuming that the forecast error of inflation follows a white noise process and test the parity condition although the RIP hypothesis is based on ex ante and not ex post real rates. Since market integration may be falsely rejected due to biased expectations, Cumby & Mishkin (1986) and Mishkin (1981, 1984b) measure ex ante real interest rates. Besides, Dutton (1993) constructs inflation rates using a price index based on traded goods because perfect goods market arbitrage may be violated for nontraded goods. Comparing her result to those obtained from CPI and WPI data the RIP is rejected less often whenever the traded goods measure is applied - except for the countries with capital mobility restrictions like France and Japan. However, considering different inflation measures and ex ante interest rates affect results only slightly; in early studies short run RIP can be clearly rejected, contradicting the integration hypothesis.[1]

However, none of the studies account for the time series properties of the underlying data. If inflation and nominal interest rates are indeed nonstationary, t-test results may be misleading because the coefficient standard errors are not consistently estimated (cf. Rose 1988; Walsh 1987; Garcia and Perron 1996, among others). Moreover, if one does not account for cointegration among nominal rates and inflation, model misspecification is an issue. Thus, recent research focuses on time series techniques such as unit root and cointegration testing and Granger causality tests.

In a cross-country study Goodwin & Grennes (1994) perform unit root and cointegration tests on ex post real rates for a set of 10 industrialized countries over the sample period from 1957-1987. The high persistence and volatility of real interest rates observed in the 1980s imply nonstationarity in the real interest rate data, but the RIP may be reflected in a cointegrating relationship. The authors confirm this hypothesis from bivariate analyses with especially strong support for Canada and the UK. Within a multivariate framework, the number of cointegration relations reflects a high degree of financial market integration. For five major

[1] There is only limited support for convergence looking at longer maturity bond rates for Canada and the UK (see, e.g., Jorion 1996; Fujii and Chinn 2000).

economies Goodwin & Grennes (1994) find four cointegrating relations and thereby confirm a high degree of market integration. Similarly, Awad & Goodwin (1998) find evidence for a dynamic linkage between weekly real interest rates among G10 countries within a sample period from 1976-1994 in the short and in the long run. In half of the considered country pairs cointegration testing points to strong financial market integration and a stable long run relation between the real rates. However, the parity condition does not hold as the intercept term is significant. The authors relate this to transaction costs. Granger causality tests show that the US rates do not respond to other rate changes reflecting the dominance of the US in world financial markets and the fact that the US was among the first to loosen restrictions to capital mobility. In contrast, shocks to the US yield immediate and long lasting responses in Canada and the UK, and less so in Germany and France. More recently Felmingham & Cooray (2008) confirm their findings within a vector error correction model for the G7 countries from 1970-2003 and find a high degree of interdependence in the long run and an increase in integration over the sample period. Also, Granger tests provide strong exogeneity evidence for the US and Germany.

Recent research accounts for nonlinearities and structural breaks when analyzing real rate convergence within a time series framework. Among others, Ferreira & Leon-Ledesma (2007) and Mancuso, Goodwin & Grennes (2003) conduct unit root tests within threshold autoregression models (TAR) for several OECD countries and find nonlinear mean reversion for the real interest rate differential to the US, the latter study confirming real rate covergence especially for Canada and the UK, and in part for Germany and Japan. Besides, Arghyrou, Gregoriou & Kontonikas (2009) find stationary differentials for 21 out of EU25 countries against the EMU average 1996-2005 allowing for two endogenous structural breaks in the real rate differentials. Research on G7 countries find similar evidence for the EU-US relationship, and in between 1974 and 1995 bilateral convergence between the US and the other G7 countries is evident (Fountas and Wu 2000, 1999).

To sum up, for datasets comprising only the recent years real interest rate convergence is often confirmed in time series analysis but over a broad sample RIP is rejected to hold for most specifications. However, the shortcomings of the time series approach on interest rate convergence are obvious: Univariate unit root tests have low power against a stationary alternative, and Granger causality tests may suffer from an omitted variable bias, i.e., a third variable might drive real interest rates at different lags.

Besides, evidence on real rate convergence is still inconclusive and an interpretation in terms of missing financial market integration may be misleading. Various factors must be considered as the potential causes of persistent deviations from the RIP may be multifold. First, we do observe transaction costs (see, e.g., Goodwin and Grennes 1994; Awad and Goodwin 1998). Second, nontraded goods entering the price index measure may bias the result. Third and most important, violations of the UIP cause serious RIP violations (Cung and Crowder 2004; Ferreira 2004). Indeed, there is evidence for a nonstationary risk premium. However, central bank monetary policy reactions to exchange rate variations may also drive nonstationarity of short-term interest rate differentials indicating that interest rates may even be I(2) (cf. Matros and Weber 2011).

With respect to the results of Matros & Weber (2011) and the methodological shortcomings of the regression-based tests and time series analysis, we propose to analyze financial market integration by looking at real interest rate convergence across countries using a different method, the dynamic latent factor model. Doing so, we extend the current literature twofold. First, besides some studies on panel unit root tests (see, e.g., Wu and Chen 1998)[2], recent literature does not analyze joint real rate convergence in a multi-

[2]Wu & Chen (1998) use a battery of panel unit root tests on CPI-based annualized ex post real interest rate data for several OECD countries from 1979 to 1996. Such panel tests have greater power than single equation augmented Dickey-Fuller tests. Joint testing supports stationary interest rate differentials suggesting increased market integration. The results are robust to the chosen

country framework, but just looks at country pairs. We aim to close this gap. Second, we are able to extract common factors in the interest rates across countries. If these stochastic factors can be reduced over time, this would be a strong indication for convergence among the considered real rate data. To the best of our knowledge we are the first who analyze multi-country real interest rate convergence within this modeling framework. Due to the evidence of a random walk behavior in the real rate, we decompose the real interest rates across several developed countries into latent factors, i.e., common permanent and transitory components, and an idiosyncratic error term for each country. The permanent components capture the random walk component in the real interest rates whereas the transitory components are modeled as autoregressive (AR) processes.[3] Thus, we capture short and long run dynamics in the real interest rate convergence process.

Beyond missing market integration, theory suggests several possible explanations of persistent deviations from the RIP (briefly outlined above). Although it is good practice to refrain from interpreting the factors we can cautiously associate the unobserved components with these different explanations of the failure of RIP to hold empirically either in the short or in the long run. First, Chung & Crowder (2004) and Ferreira (2004) confirm the theory that a violation of the UIP is the most likely candidate to cause serious RIP violations. Studies concerning the UIP typically find a nonstationary risk premium. This may result from deviations from the risk neutrality assumption, heterogeneous agents or heterogeneous information in financial markets, or deviations from the rational expectation assumption. Concerning our modeling framework, a transitory or permanent time-varying risk premium due to heterogeneous agents may be reflected in one common component. Second, PPP does not hold at a short horizon. Nontraded goods that enter the price index measure used in empirical studies may bias the result. However, long horizon price equalization is verified in the empirical literature. Thus, in our model transaction costs are well captured in the common transitory component or in an idiosyncratic error term. In addition, there may be a strong downward trend in transaction costs in the recent past most likely captured in the permanent components' movements. Third, the permanent and transitory components may reflect policy changes in large countries that influence the other countries. Thus, since the different reasons for a RIP failure are difficult to model and, more important, to measure, it is straightforward to resort to a factor model with common factors.

To analyze changing convergence patterns over time we take a close look at modeling the variances of the components as time-varying. We directly measure convergence over time by using a time dependent logistic function. This is in line with recent research. For example Luginbuhl & Koopman (2004) consider a dynamic latent factor model to analyze GDP convergence among European countries. The authors differentiate between convergence in the trend and cycle series, and in the variances by introducing a smooth transition term in form of a logistic function that depends on time. In line with this, Pozzi & Wolswijk (2012) analyze if government bond risk premia are driven by common factors and/or by a time-varying country-specific stochastic factor. Time variation is again modeled via a logistic function and the authors indeed find that the exposure to common international risk factors has converged across countries, and towards zero for the idiosyncratic component in the risk premia.

Our estimation results can be briefly summarized as follows. Taken together, four permanent and four transitory components capture the real interest rate dynamics among the six industrialized countries. We regularly observe decreasing common component variances over time, in part even converging to values close to zero. Thus, this indicates a possible reduction in the number of stochastic components over time, which is indeed strong support for the convergence hypothesis. The transition functions show a smooth transition between variance regimes. We interpret this in favor of an overall slow and still ongoing con-

base country, except when using the US as base country. This supports the result of Gagnon & Unferth (1995) who show that the world real interest rate differential is just serially correlated using US data, but not for the other tested nine OECD countries.

[3]Similarly, Neely & Rapach (2011) study inflation convergence in OECD countries since the 1950s using a dynamic latent factor model with unobserved world, regional and idiosyncratic autoregressive factors.

vergence process among real interest rates. After rapid convergence during the late 1970s and 1980s, we observe slower transition since the mid 1990s. This may reflect the fact that financial markets have already been highly integrated in the 1990s. The speed and timing of the convergence process is similar among the permanent and transitory states.

The remainder of the chapter is structured as follows. The dataset is described in section 2, the modeling framework is established in detail in section 3. Secion 4 provides the estimation results together with an interpretation, and section 5 conludes.

2 The dataset

Our analysis is based on monthly data over a sample period from 1975M1 to 2011M3 for six OECD countries. Nominal interest rates, i.e., short-term money market rates (federal fund rates) are taken from the IMF International Financial Statistics (IFS). We employ the widely used CPI data to compare our results to the findings of former studies. The CPI index is provided by the OECD Main Economic Indicators.

We follow other studies (Mishkin 1984a; Cumby and Obstfeld 1984) and base our analysis on the ex post real interest rate. The major difficulty is that expected inflation and thus ex ante interest rates are not observable. Indeed, under the rational expectations assumption, expected inflation is just given by realized inflation plus some random error term. If the estimated error term is not white noise, by constructing ex ante data within a regression-based framework we might run into an errors-in-variable problem. Therefore, we use price data from the last month of the previous quarter to the current quarter to construct three-month realized inflation rates as this corresponds to the maturity considered in the nominal rates.[4]

We substract realized inflation from the nominal rates to construct the ex post real rates used in our analysis on the following six G7 countries: the United States, United Kingdom, Germany, Japan, Canada and Italy. We disregard France from our analysis for two reasons. First, Granger causality test results from Awad & Goodwin (1998) suggest that France not Germany seems to have a leading role in the EMS - French rates do not respond to shocks to other European rates. However, France was never dominant in international financial markets. The authors relate their findings to the capital mobility barriers that were loosened late in 1986 and to the observed slow adjustment to long run equilibrium rates, i.e., the correlation might be spurious. Second, a more practical issue arises because the money market rate is not available for France over this time horizon using the IMF IFS database.

3 The modeling framework

We consider a dynamic latent factor model similar to the one proposed in Luginbuhl & Koopman (2004). The observation equation in our model describes the decomposition of the real interest rate r_t for P countries into two state vectors. The first $(K_1 \times 1)$ state vector μ_t contains the K_1 permanent factors whereas the second describes the transitory factors ε_t being of dimension $(K_2 \times 1)$.

$$r_t = c + A\mu_t + B\varepsilon_t + \nu_t, \quad t = \ldots, T. \tag{3.1}$$

c is a $(P \times 1)$ constant term with nonzero parameter values for the last $(P - K_1)$ entries and zero otherwise. In addition A and B are $(P \times K_1)$ and $(P \times K_2)$ parameter matrices for the permanent and transitory factors, respectively, and ν_t is a vector of iid measurement errors of dimension $(P \times 1)$ with mean zero and time-invariant variance-covariance matrix $R_{(P \times P)}$, with $\mathrm{E}[\nu_{pt}\nu_{qs}] = 0,\ \forall\ p \neq q,\ p, q = 1, \ldots, P,$

[4] We also conducted the analysis on real interest rates constructed with the three-month ahead inflation rates. The results from this forward looking approach are qualitatively similar and available on request.

and $t \neq s,\ t,s = 1,\ldots,T$. The factors are modeled as follows. We model the permanent components as random walks. The vector of permanent components can be written as

$$\mu_t = \mu_{t-1} + \eta_t, \quad \eta_t \sim \text{iid}(0, Q_{\eta t}) \tag{3.2}$$

with $Q_{\eta t} = \text{diag}(d_{\eta,1,t},\ldots,d_{\eta,K_1,t})$. We assume that $\text{E}[\eta_{k_1,t}\eta_{l_1,s}] = 0,\ \forall k_1 \neq l_1,\ k_1,l_1 = 1,\ldots,K_1$ and $\forall t \neq s,\ t,s = 1,\ldots,T$. The transitory components are modeled as stationary AR(1) processes

$$\varepsilon_t = \alpha\varepsilon_{t-1} + \zeta_t, \quad \zeta_t \sim \text{iid}(0, Q_{\zeta t}) \tag{3.3}$$

with $\alpha = \text{diag}(\alpha_1,\ldots,\alpha_{K_2})$, $Q_{\zeta t} = \text{diag}(d_{\zeta,1,t},\ldots,d_{\zeta,K_1,t})$, and again $\text{E}[\zeta_{k_2,t}\zeta_{l_2,s}] = 0,\ \forall k_2 \neq l_2,\ k_2,l_2 = 1,\ldots,K_2$ and $\forall t \neq s,\ t,s = 1,\ldots,T$. The AR lag order for the transitory factors is chosen as follows. Since the differences between the data and the nonstationary factors reflect short run changes in the real interest rates, they may be well approximated by the first differences of the real rates. Thus, we fit a VAR for the first differences of our real interest rate data. The optimal lag length for the VAR following the Schwarz criterion is one. Thus, for the four transitory factors in our model we suggest a lag order of 1 and thus implement AR(1) processes.[5]

The permanent and transitory factor innovations are uncorrelated at all leads and lags. The factor variances are time-varying such that

$$d_{\eta,k_1,t} = \kappa_{\eta t}\sigma^2_{\eta,k_1}, \quad \kappa_{\eta t} = \exp[\xi_\eta(t-\tau_\eta)]/\{1+\exp[\xi_\eta(t-\tau_\eta)]\}, \tag{3.4}$$

$$d_{\zeta,k_2,t} = \kappa_{\zeta t}\sigma^2_{\zeta,k_2}, \quad \kappa_{\zeta t} = \exp[\xi_\zeta(t-\tau_\zeta)]/\{1+\exp[\xi_\zeta(t-\tau_\zeta)]\}, \tag{3.5}$$

$t = 1\ldots,T$. We consider the possibility of a variance reduction over time by multiplying $0 \leq \kappa_{\eta t} \leq 1$ and $0 \leq \kappa_{\zeta t} \leq 1,\ t = 1,\ldots,T$ with the constants σ^2_{η,k_1} and σ^2_{ζ,k_2}, respectively. Thereby, we assume two common convergence functions, one among all permanent and one among all transitory factors. We allow for differently shaped transition functions between the transitory and permanent variances, and we allow for different σs for each single factor.

In detail, κ_η and κ_ζ are described as logistic functions, for which the shape depends on realized parameter values for ξ_η and τ_η, and ξ_ζ and τ_ζ, resp., and a time trend t. Disregarding the subscripts, the transition function behavior can be described as follows. ξ denotes the slope parameter of the transition function and thus measures the rate of convergence. High absolute values for ξ indicate a steep transition from one regime to another, whereas low values imply a rather smooth convergence. For $\xi = 0$ there is no convergence in the corresponding common component. In a sample of size T, $\xi < 0$ implies that κ goes to 0 for $t >> \tau$ and that κ goes to 1 for $t << \tau$. The second parameter $1 \leq \tau \leq T$ can be interpreted as location parameter, determining when a regime change sets in. We assume two different regimes. Using the time trend as a transition variable we further assume that only one transition towards equal interest rates takes place within the chosen time period. If we observe convergence in real rates, the diagonal elements of the variance-covariance matrices Q_η and Q_ζ should converge to zero for the corresponding components.

The optimal number of latent factors to include is of course unknown. Intuitively, we start with imposing a large number of unobserved factors for the first point in time. The number of stochastic factors may be further reduced over time through a decrease in the variances (as described above). More important, the choice of the number of factors may be based on different criteria. For an overview on different approaches we refer to Breitung & Eickmeier (2006). We base our decision on the final number of components on the results from a principal component analysis (PCA) of the real interest rates for the first five years of

[5]Moreover, using a higher AR order would make the estimation quite cumbersome. Moving from AR(1) to AR(2) increases the number of parameters and additionally doubles the number of unobserved transitory states.

the sample. In our PCA the first four principal components already explain about 91 percent of the total variance. Since there is no overall optimal criterion for the factor choice, the model is also estimated for different numbers of transitory and permanent components, K_1 and K_2. However, the PCA and diagnostic tests performed after the estimation suggest to use $K_1 = K_2 = 4$. Detailed results are available on request.

Besides the number of factors, the empirical results also depend on the order of the countries. The country whose real interest rate is the most reasonable candidate to influence the other rates should be placed first. The country that may be the most integrated should be placed last. Based on the findings in the literature and based on country size, we decide on the following order for the $P = 6$ considered countries: The US is placed first as it takes on the leading role in the world economy. From the Granger causality tests and regression analysis presented in the introduction we know that Germany is indicated as a leading economy, especially among Europe. The United Kingdom, however, is an important financial market place. Thus, since UIP deviations are more likely to explain the RIP failure than the short run deviations of the PPP, we decide to put the UK before Germany in the subsequent analysis.[6] Japan is placed next. Finally we position Canada and Italy as the results from time series analysis cited above confirm that Canada is closely linked to the US, whereas Italy can be considered a small country, at least from a financial market perspective.

4 Empirical results

The results from maximum likelihood estimation are presented in Tables 1 and 2. The transition functions as well as the filtered components using the Kalman filter are displayed in Figures 1 and 2.[7]

Table 1: Estimation results for the state space model, factor loadings

	UK	Germany	Japan	Canada	Italy
γ_1	0.321	0.261	0.118	0.602	0.290
	(0.071)	(0.082)	(0.066)	(0.034)	(0.063)
γ_2		0.330	0.233	0.415	0.201
		(0.107)	(0.127)	(0.075)	(0.164)
γ_3			1.288	0.749	0.088
			(0.382)	(0.203)	(0.422)
γ_4				0.399	4.210
				(0.233)	(1.619)
θ_1	-0.080	-0.830	0.993	0.661	-0.431
	(0.266)	(0.493)	(0.448)	(0.130)	(0.316)
θ_2		-0.099	0.254	-0.110	0.018
		(0.045)	(0.060)	(0.066)	(0.049)
θ_3			-3.396	-2.375	-0.672
			(1.173)	(2.280)	(0.445)
θ_4				-2.128	0.155
				(1.101)	(0.114)

Note: Standard errors are given in parentheses.

The estimates for the factor loading coefficients in Table 1 are almost all significant. We find strong support for the importance of all four permanent factors, except for Italy for the third (US-UK-GER) permanent

[6] In a robustness check at the end of section 4 we change the order and obtain qualitatively and quantitatively similar results.

[7] Details on the estimation method and imposed identifying restrictions are given in the mathematical appendix.

component. As expected, the first permanent component in the upper left panel of Figure 1 maps the long run movements of the real interest rate in the US. Moreover, all the remaining components are decreasing over time, as well as the corresponding volatilities.

For the transitory component, the US component does not significantly contribute to the UK and the German real rate, and the coefficient of the second transitory component (US-UK) is insignificant for Italy. The latter may be explained by the fact that indeed Germany is the leading economy in Europe and Italy's short run movements in the real interest rate may be most likely synchronized to German rates. (The case is different for the permanent components, which might indicate that Italy is exposed to permanent changes in the world leading economies.) The transitory components are presented in the lower panels in Figure 1. Consequently, the first component in the upper left reflects the transitory movements of US real interest rates. The other three panels show strong support for the convergence hypothesis in the transitory components of real interest rates among countries; volatility decreases over time until the mid 1990s and stays low afterwards.

The estimated AR parameters presented in Table 2 are all well below unity. The imposed stationarity restriction is not binding. The first component is the most persistent one with $\tilde{\alpha}_1 = 0.957$, the other estimates are in a range of about 0.59 to 0.73. We can conclude that the transitory parts in real interest rates are of relevance.

The slope and location parameter of the two transition functions are quite similar (cf. Table 2 and Figure 2, left panel). In general, the variances are rapidly converging to the lower regime in the late 1970s and 1980s, and continue to converge at a slower pace from the mid 1990s onwards. This may reflect the fact that financial markets have been well integrated since the 1990s. The transition is very smooth, $\tilde{\xi}_\eta = -0.011$ and $\tilde{\xi}_\zeta = -0.012$, and the location parameter $\tilde{\tau}_\eta$ is only slightly higher than $\tilde{\tau}_\zeta$, suggesting a later transition to the low variance regime for the permanent relative to the transitory variances. This is confirmed in the right panel of Figure 2. The first difference plots visualize a rapid transition especially in the beginning of the 1980s. Since 2000 the convergence process becomes remarkably slower. Moreover, whereas until 1990 there is faster convergence in the variances of the transitory relative to the permanent components, the opposite holds for the period afterwards. Diagnostic test results are provided in Table 3. Ljung-Box test results suggest remaining autocorrelation in the estimated prediction error, except for certain lags for the UK and Canada. We do not find evidence for heteroskedasticity in the prediction error.

Compared to the literature on real interest rate convergence outlined in the introductory section we emphasize two striking differences to our results from the multi-country analysis. In contrast to Granger causality test results in the time series literature, a salient result in our analysis is that also Germany's real rates are described by permanent and transitory factors based on US and UK data. Moreover, in our setting we find no significant influence of the transitory component of the US real rates on UK real interest rates. In line with the results from panel unit root tests, we demonstrate that indeed interest rate convergence took place during the last four decades. The low estimates for the variances in recent years indicate that the interest rates may differ only with respect to some constant that might capture economic issues beyond missing market integration, e.g., transaction costs or risk aversion.

Robustness checks

We interpret the slowdown in the convergence speed from the mid 1990s onwards as a sign for highly integrated financial markets. However, the result may be driven by increased interest rate spreads during the global financial crisis period. Moreover, the rapid adjustment during the early years may simply reflect the overall decreasing volatility in economic data during the so-called great moderation. Besides, up to a certain degree the specific country order may influence the results.

Table 2: Estimation results for the state space model

Constant term					
c_1	c_2				
0.113	5.741				
(0.395)	(0.988)				
AR parameters transitory factors					
α_1	α_2	α_3	α_4		
0.957	0.595	0.649	0.725		
(0.017)	(0.077)	(0.079)	(0.036)		
Transition function parameters					
ξ_η	ξ_ζ	τ_η	τ_ζ		
-0.011	-0.012	91.308	82.352		
(0.001)	(0.000)	(4.997)	(3.242)		
Variance-covariance matrix: permanent factors					
$\sigma^2_{\eta,1}$	$\sigma^2_{\eta,2}$	$\sigma^2_{\eta,3}$	$\sigma^2_{\eta,4}$		
0.413	1.518	0.266	0.073		
(0.064)	(0.445)	(0.109)	(0.057)		
Variance-covariance matrix: transitory factors					
$\sigma^2_{\zeta,1}$	$\sigma^2_{\zeta,2}$	$\sigma^2_{\zeta,3}$	$\sigma^2_{\zeta,4}$		
0.075	4.431	0.093	0.604		
(0.064)	(0.697)	(0.069)	(0.473)		
Variance-covariance matrix: idiosyncratic error term					
US	UK	GER	JP	CAN	IT
$\sigma^2_{\nu,1}$	$\sigma^2_{\nu,2}$	$\sigma^2_{\nu,3}$	$\sigma^2_{\nu,4}$	$\sigma^2_{\nu,5}$	$\sigma_{\nu,6}$
3.681e-7	0.114	0.055	0.008	1.369e-4	1.166e-5
(8.694e-9)	(0.034)	(0.009)	(0.009)	(9.789e-6)	(1.278e-6)
ln L	AIC				
-1833.164	8.649				

Note: Standard errors are given in parentheses. ln L denotes the value of the log-likelihood function and AIC is the Akaike information criterion.

Table 3: p-values of diagnostic tests for the estimated model in Tables 1 and 2

	US	UK	Germany	Japan	Canada	Italy
Country-specific Ljung-Box test for autocorrelation, H_0 : no						
1 lag	0.000	0.292	0.001	0.004	0.868	0.005
6 lags	0.000	0.000	0.000	0.000	0.058	0.018
12 lags	0.000	0.000	0.000	0.000	0.141	0.052
Country-specific test for heteroskedasticity, H_0 : no						
	1.000	0.268	0.197	0.220	0.914	0.999

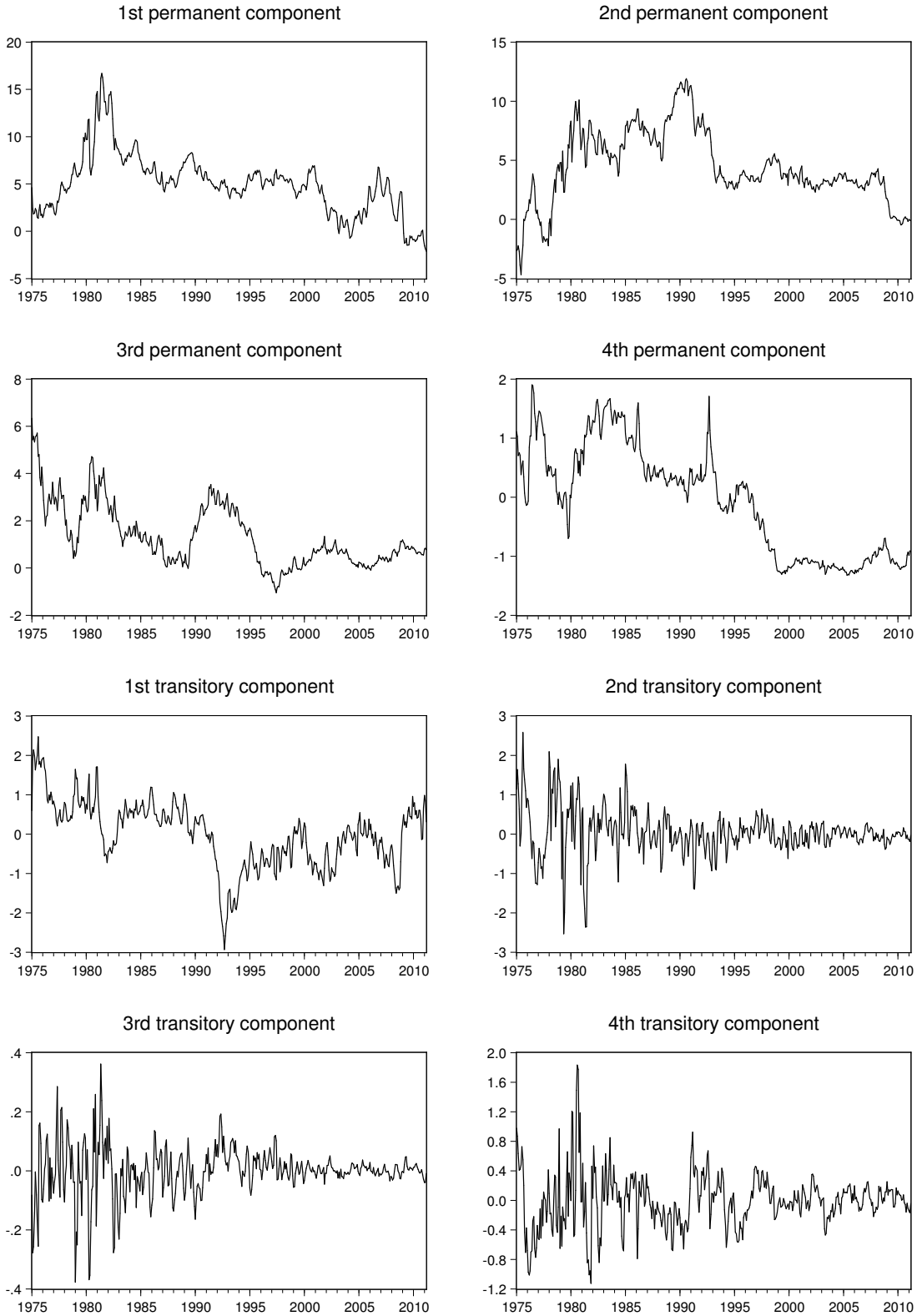

Figure 1: Permanent and transitory components for the state space model in Tables 1 and 2, The four upper panels: permanent components $\mu_{k_1},\ k_1 = 1,\ldots,4$; the four lower panels: transitory components $\varepsilon_{k_2},\ k_2 = 1,\ldots,4$.

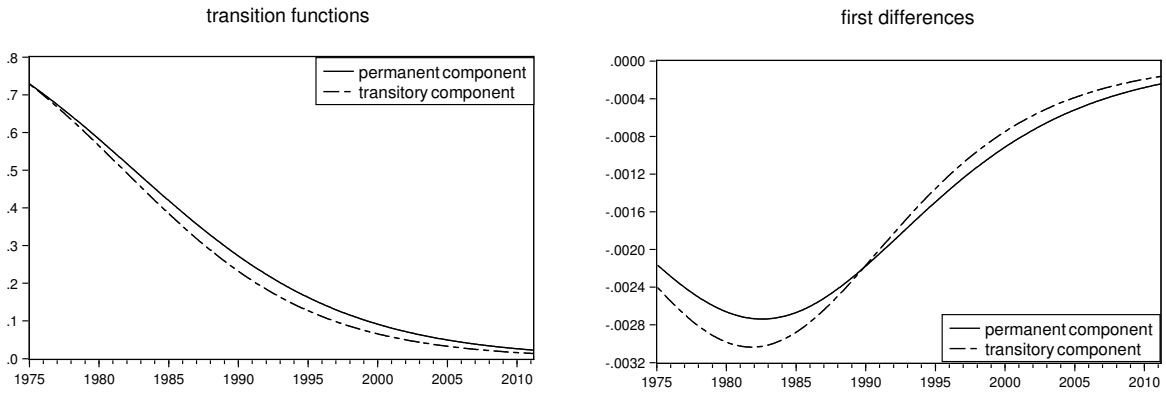

Figure 2: Left panel: transition function κ_η (time-varying part of the variances of the permanent components, solid line) and transition function κ_ζ (time-varying part of the variances of the transitory components, dashed line). Right panel: corresponding first differences of the transition functions.

Thus, to check the robustness of our results we discuss two changes compared to the baseline specification, First, we reestimate the model for different subperiods. Compared to the full sample result from 1975M1 to 2011M3 presented above, the first subperiod includes only the years for 1975M1 to 2007M1, the second and third begin in 1985M1 and last until 2005M12 and 2011M3, respectively. Second, since the country ordering cannot be based on objective criteria, we change the order and put Germany in second place. Although the UK is an important financial hub, Germany is a leading economy with high export sales and has been among the first countries that abandoned capital controls in 1973.

The two left panels of Figure 3 show the resulting transition functions for the permanent and transitory components' variances for the different sample sizes. For the time period 1975M1 to 2006M12 (short-dashed line) our results remain qualitatively the same. Although the speed of adjustment increases in the beginning, the slowdown in the speed of adjustment around 2000 and variance estimates close to zero are still obtained using the shorter subsample. Starting in 1985, the convergence speed even increases considering the dynamics for the permanent components and only decreases in the transitory components' variance function including the crisis years (long-dashed and long-short dasehd lines). However, besides small changes in the patterns due to the recent crisis, the outcome is not driven by the chosen sample size, and convergence still takes place in the permanent as well as the transitoy component variances. All these results also obtain for the two different country orders (right panels of Figure 3).

The estimation results for the full sample and the revised country order are presented in Tables 4 and 5. In contrast to the results presented above (Tables 1 and 2), the factor loadings of the third permanent factor (US-GER-UK) are not significant for Japan and Italy. Moreover, the loadings of the third (and fourth) transitory components are not significant for Japan, Canada and Italy. In addition, the estimated US transitory factor loading $\tilde{\theta}_{13}$ is not significant for the UK data. The filtered components as well as the transition functions look very similar to those shown in Figures 1 and 2. Figures 4 and 5 illustrate the differences between the components including the different subsamples. (The corresponding tabulated estimation results as well as diagnostic tests for the three subsamples are not shown to conserve space.) The overall impression is that the permanent factors look very similar, except for small shifts in the level. They almost all coverge to zero, independent of the country order and the chosen sample size. As expected, the second and third permanent factors' shapes are inverted for the new country order.

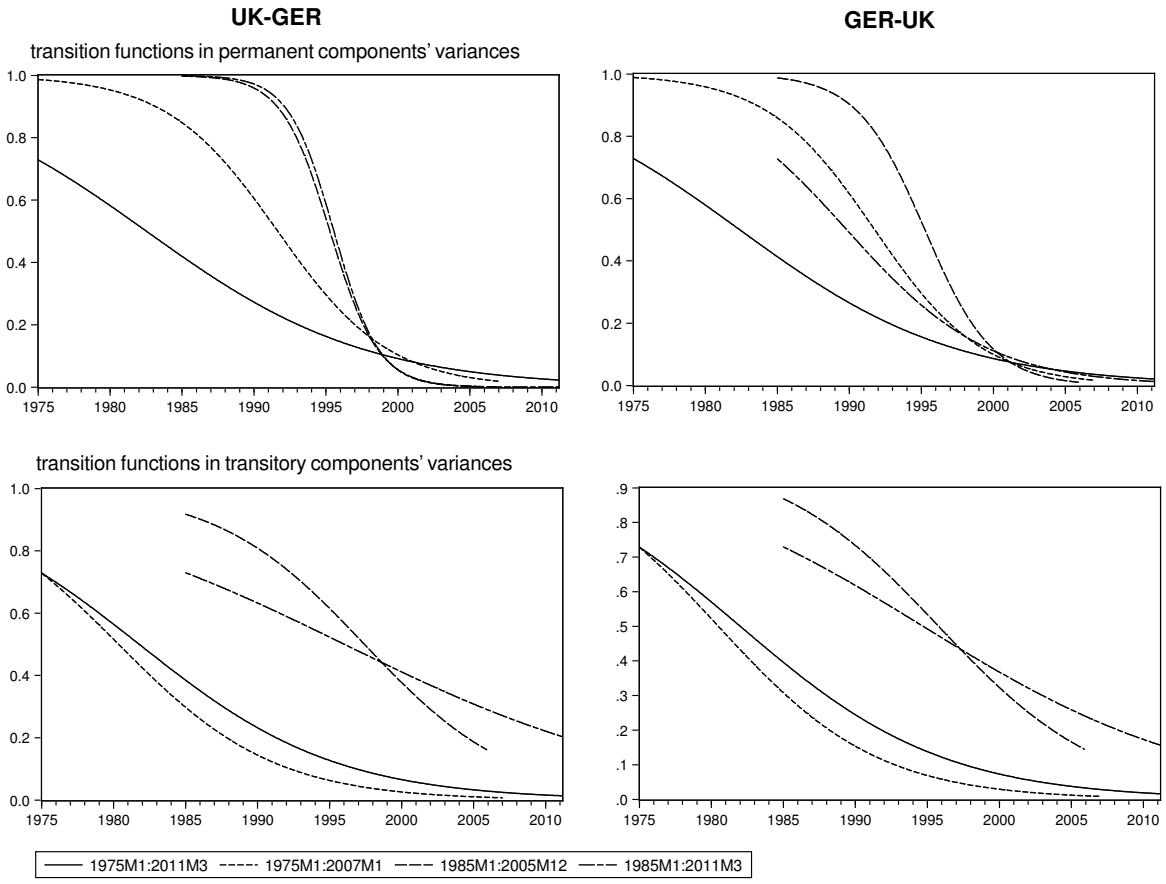

Figure 3: Left panels: transition functions for the permanent (upper panel) and transitory components (lower panel) for all subsamples using the baseline country order. Right panels: transition functions for the permanent (upper panel) and transitory components (lower panel) for all subsamples for the model with the UK ordered before Germany. Solid lines always indicate the full sample results.

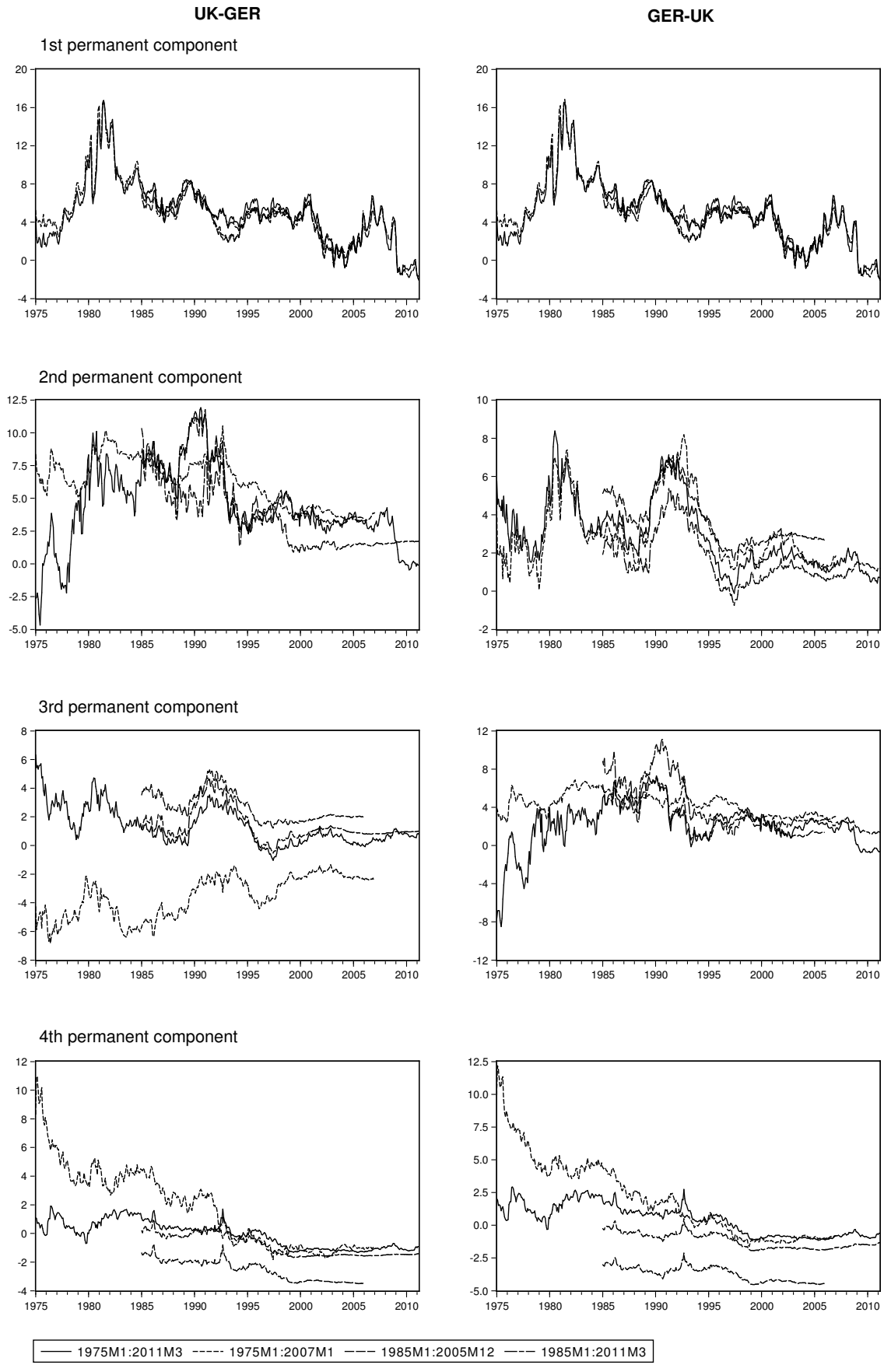

Figure 4: Left four panels: permanent components ($\mu_{k_1},\ k_1 = 1,\ldots,4$) for the state space model in Tables 1 and 2 and the results for the subperiods. Right four panels: permanent components for the state space model in Tables 4 and 5 (model with Germany ordered first), and the results for the subperiods.

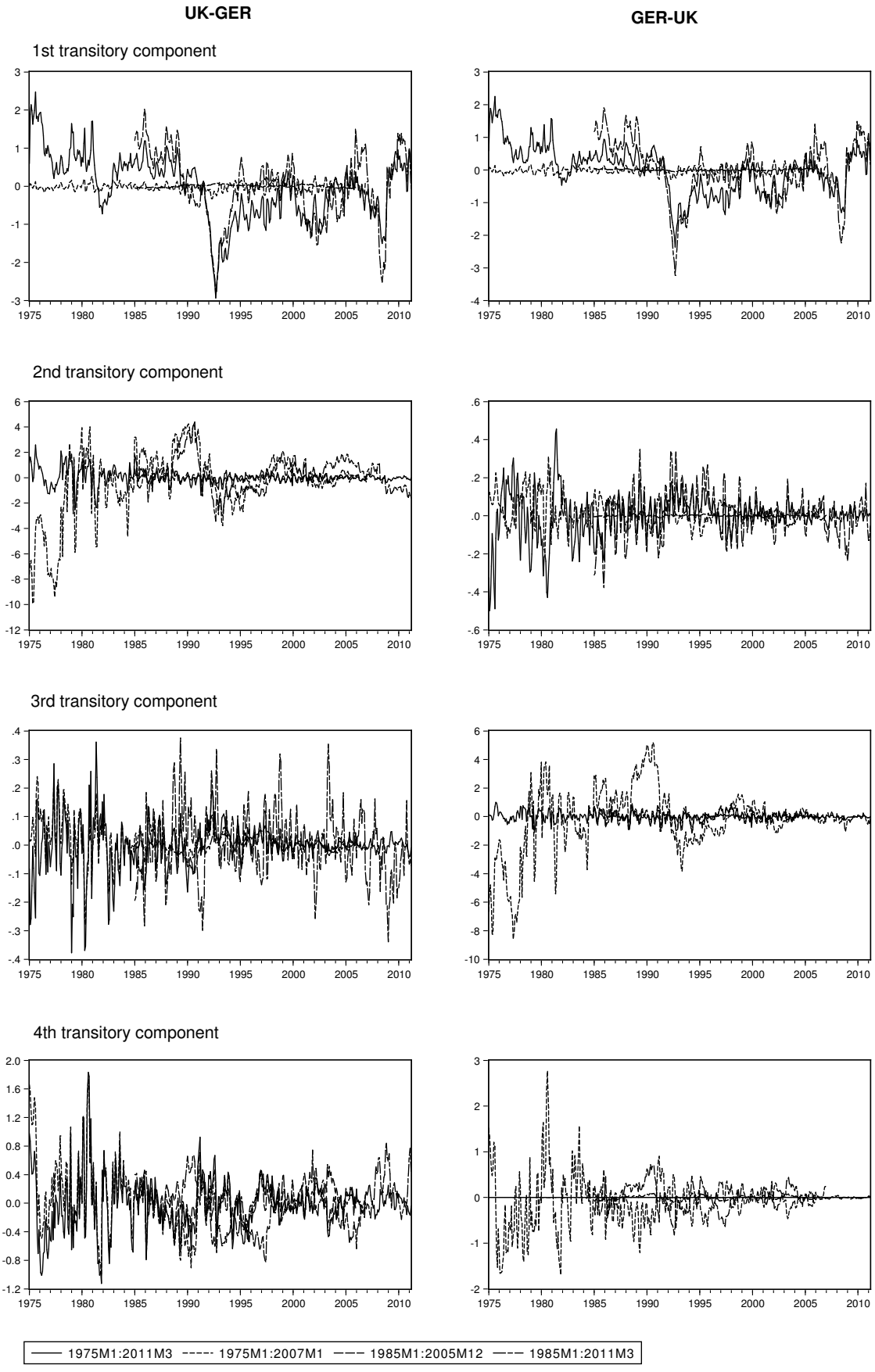

Figure 5: Left four panels: transitory components ($\varepsilon_{k_1},\ k_1 = 1, \ldots, 4$) for the state space model in Tables 1 and 2 and the results for the subperiods. Right four panels: transitory components for the state space model in Tables 4 and 5 (model with Germany ordered first), and the results for the subperiods.

Table 4: Estimation results for the state space model, factor loadings - different country order

	Germany	UK	Japan	Canada	Italy
γ_1	0.238	0.319	0.143	0.603	0.277
	(0.097)	(0.077)	(0.080)	(0.037)	(0.072)
γ_2		0.792	0.846	0.638	0.237
		(0.303)	(0.217)	(0.122)	(0.160)
γ_3			−0.174	0.227	0.193
			(0.200)	(0.097)	(0.180)
γ_4				0.513	3.308
				(0.210)	(1.622)
θ_1	−0.926	−0.185	1.082	0.676	0.518
	(0.704)	(0.406)	(0.635)	(0.159)	(0.491)
θ_2		−4.563	−5.459	1.016	−0.717
		(2.181)	(1.738)	(0.558)	(0.467)
θ_3			−0.104	−1.959e-4	−0.061
			(0.186)	(0.086)	(0.070)
θ_4				810.560	−4.093
				(813.170)	(29.206)

Note: Standard errors are given in parentheses.

5 Conclusion

We model real interest rate convergence within a multi-country framework by means of a dynamic latent factor model. We use interest rate data from 1975M1-2011M3 for six industrialized countries. We decompose the real rates into permanent and transitory factors. The variances are modeled as time-varying processes allowing for one endogenous transition from a high variance regime at the beginning of the sample towards a low variance regime. We find a decline in the variances for the permanent as well as for the transitory latent factors, in part decreasing to values close to zero. We interpret this in favor of real interest rate convergence among the sample economies.

Interestingly, although there was a fixed point in time in which each analyzed country relaxed its capital controls in the 1970s and 1980s, the transition to a low variance regime is smooth. After a rapid convergence process in the late 1970s and 1980s, the transition towards the small variance regime has slowed down since the mid 1990s.

Table 5: Estimation results for the state space model - different country order

Constant term					
c_1	c_2				
0.322	3.925				
(0.355)	(3.782)				
AR parameters transitory factors					
α_1	α_2	α_3	α_4		
0.958	0.701	0.423	0.706		
(0.019)	(0.051)	(0.119)	(0.047)		
Transition function parameters					
ξ_η	ξ_ζ	τ_η	τ_ζ		
-0.011	-0.012	89.649	85.035		
(0.001)	(0.001)	(5.134)	(3.897)		
Variance-covariance matrix: permanent factors					
$\sigma^2_{\eta,1}$	$\sigma^2_{\eta,2}$	$\sigma^2_{\eta,3}$	$\sigma^2_{\eta,4}$		
0.428	0.487	1.253	0.122		
(0.071)	(0.084)	(0.462)	(0.118)		
Variance-covariance matrix: transitory factors					
$\sigma^2_{\zeta,1}$	$\sigma^2_{\zeta,2}$	$\sigma^2_{\zeta,3}$	$\sigma^2_{\zeta,4}$		
0.061	0.070	2.727	4.953e-6		
(0.072)	(0.043)	(1.028)	(9.656e-6)		
Variance-covariance matrix: idiosyncratic error term					
US	UK	GER	JP	CAN	IT
$\sigma^2_{\nu,1}$	$\sigma^2_{\nu,2}$	$\sigma^2_{\nu,3}$	$\sigma^2_{\nu,4}$	$\sigma^2_{\nu,5}$	$\sigma_{\nu,6}$
1.213e-147	0.060	0.101	1.322e-94	6.501e-31	1.316e-137
(4.017e-145)	(0.010)	(0.034)	(2.776e-92)	(2.145e-29)	(4.018e-135)
$\ln$ L	AIC				
-1834.824	8.657				

Note: Standard errors are given in parentheses. $\ln$ L denotes the value of the log-likelihood function and AIC is the Akaike information criterion.

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6 Mathematical appendix

The model in Eqs. (3.1)-(3.3) can be written in state space matrix notation. The observation equation is then given by

$$r_t = c + \left[\Gamma \vdots \Theta\right] \left[\frac{\mu_t}{\varepsilon_t}\right] + \nu_t, \quad \nu_t \sim \text{iid}(0, R) \tag{6.1}$$

with R being the error variance-covariance matrix. The state equation is denoted by

$$\left[\frac{\mu_t}{\varepsilon_t}\right] = \left[\begin{array}{c:c} I_{(K_1 \times K_1)} & 0_{(K_1 \times K_2)} \\ \hdashline 0_{(K_2 \times K_1)} & \alpha_{(K_2 \times K_2)} \end{array}\right] + \left[\frac{\mu_{t-1}}{\varepsilon_{t-1}}\right] + \left[\frac{\eta_t}{\zeta_t}\right] \tag{6.2}$$

and the state error variance-covariance matrix can be written as

$$Q_t = \left[\begin{array}{c:c} Q_{K1,t} & 0_{(K_1 \times K_2)} \\ \hdashline 0_{(K_2 \times K_1)} & Q_{K2,t} \end{array}\right]. \tag{6.3}$$

In detail,

$$\begin{pmatrix} r_{1t} \\ r_{2t} \\ \vdots \\ \vdots \\ \vdots \\ \vdots \\ \vdots \\ r_{Pt} \end{pmatrix} = \begin{pmatrix} 1 & 0 & \cdots & 0 & 1 & 0 & \cdots & 0 \\ \gamma_{21} & \ddots & \ddots & \vdots & \theta_{21} & \ddots & \ddots & \vdots \\ \gamma_{31} & \gamma_{32} & \ddots & 0 & \theta_{31} & \theta_{32} & \ddots & 0 \\ \vdots & & \ddots & 1 & \vdots & & \ddots & 1 \\ \vdots & & & \gamma_{K_1+1,K_1} & \vdots & & & \theta_{K_2+1,K_2} \\ \vdots & & & \vdots & \vdots & & & \vdots \\ \gamma_{P1} & \gamma_{P2} & \cdots & \gamma_{P,K_1} & \theta_{P1} & \theta_{P2} & \cdots & \theta_{P,K_2} \end{pmatrix} \begin{pmatrix} \mu_{1t} \\ \mu_{2t} \\ \vdots \\ \mu_{K_1,t} \\ \varepsilon_{1t} \\ \varepsilon_{2t} \\ \vdots \\ \varepsilon_{K_2,t} \end{pmatrix}$$

$$+ \begin{pmatrix} 0 \\ \vdots \\ \vdots \\ 0 \\ c_1 \\ \vdots \\ \vdots \\ c_{P-K_1} \end{pmatrix} + \begin{pmatrix} \nu_{1t} \\ \nu_{2t} \\ \vdots \\ \vdots \\ \vdots \\ \vdots \\ \vdots \\ \nu_{Pt} \end{pmatrix} \quad \text{and } R = \begin{pmatrix} \sigma^2_{\nu 1} & 0 & \cdots & 0 & 0 \\ 0 & \sigma^2_{\nu 2} & & & 0 \\ \vdots & & \ddots & & \vdots \\ 0 & 0 & \cdots & \cdots & \sigma^2_{\nu P} \end{pmatrix}, \tag{6.4}$$

$$\begin{pmatrix} \mu_{1t} \\ \mu_{2t} \\ \vdots \\ \mu_{K_1,t} \\ \varepsilon_{1t} \\ \varepsilon_{2t} \\ \vdots \\ \varepsilon_{K_2,t} \end{pmatrix} = \begin{pmatrix} 1 & 0 & \cdots & 0 & 0 & \cdots & \cdots & 0 \\ 0 & 1 & \ddots & \vdots & \vdots & \ddots & & \vdots \\ \vdots & \ddots & \ddots & 0 & \vdots & & \ddots & \vdots \\ 0 & \cdots & 0 & 1 & 0 & \cdots & \cdots & 0 \\ 0 & \cdots & \cdots & 0 & \alpha_1 & 0 & \cdots & 0 \\ \vdots & \ddots & & \vdots & 0 & \alpha_2 & \ddots & \vdots \\ \vdots & & \ddots & \vdots & \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & \cdots & 0 & 0 & \cdots & 0 & \alpha_{K_2} \end{pmatrix} \begin{pmatrix} \mu_{1,t-1} \\ \mu_{2,t-1} \\ \vdots \\ \mu_{K_1,t-1} \\ \varepsilon_{1,t-1} \\ \varepsilon_{2,t-1} \\ \vdots \\ \varepsilon_{K_2,t-1} \end{pmatrix} + \begin{pmatrix} \eta_{1t} \\ \eta_{2t} \\ \vdots \\ \eta_{K_1,t} \\ \zeta_{1t} \\ \zeta_{2t} \\ \vdots \\ \zeta_{K_2,t} \end{pmatrix}, \tag{6.5}$$

$$Q_{K1,t} = \begin{pmatrix} d_{\eta,1,t} & 0 & \dots & 0 \\ 0 & d_{\eta,2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \dots & 0 & d_{\eta,K_1,t} \end{pmatrix} \quad \text{and} \quad Q_{K2,t} = \begin{pmatrix} d_{\zeta,1,t} & 0 & \dots & 0 \\ 0 & d_{\zeta,2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \dots & 0 & d_{\zeta,K_2,t} \end{pmatrix}. \quad (6.6)$$

Our factor model is identified using a recursive structure in the factor loadings (see Eq. (6.4)). The model is estimated by maximum likelihood using the BFGS algorithm. The Kalman filter is used to obtain the states, i.e., the permanent and transitory components. Details on this approach are described in Kim & Nelson (1999).

To estimate the model we introduce the following parameter restrictions. The variances are restricted to take on only positive values. Moreover, to assure stationarity for the transitory components, the AR parameters α_{k_2}, $k_2 = 1, \dots, K_2$ are restricted to be below unity in absolute values. Most important is our identifying restriction on the slope parameter ξ. (In what follows we again skip the subscripts.) We restrict it to negative values by imposing $\xi < -1/T$. The location parameter τ is restricted such that it takes on values in between zero and T.

The starting values for the estimation process are set as follows. For the factor loadings and the unrestricted AR coefficients of the transitory factors as well as for the unrestricted variances, we randomly draw parameter starting values from a normal distribution. As it is common in the smooth transition regression literature, the choice of the starting values for the parameters ξ and τ is based on a grid search algorithm. We use two criteria, i.e., the AIC and the log-likelihood function value. The chosen starting values for our prefered model specification are $\xi_0 = -6$ and $\tau_0 = 1$, which are identical to $\xi_{restr.0} \approx -0.005$ and $\tau_{restr.0} \approx 269,300$, respectively, considering the parameter restrictions.