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Exchange and Interest Rates prior to EMU: The Case of Greece

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Exchange and Interest Rates prior to EMU: The Case of Greece*

ABSTRACT

Recently a variety of exchange and interest rate models capturing the dynamics during the transition from an exchange rate arrangement of floating rates into a currency union have been derived. While these stochastic equilibrium models in continous time are theoretically rigorous, a systematic and extensive empirical validation is still lacking. Using exchange and interest rate data collected prior to the Greek EMU-entrance on 1 January 2001 this paper tries to fill the gap between theory and real-world data. The analysis reveals that the formal models can explain many features of the Greek exchange and interest rate dynamics on the road to EMU.

JEL-classification: E43; F31; F33; C51; C52

Keywords: EMU; exchange and interest rate models; policy shifts; economic regime switching models

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1 Introduction

The dynamics of exchange and interest rates marks a topic of vital importance for many aspects in international macroeconomics and finance. Consequently, models for both economic variables abound in the literature and range from structural macromodels over microeconomic trading-models to merely empirical approaches. In a series of theoretical papers Wilfling and Maennig (2001) and Wilfling (2002, 2003a, 2003b) establish closed-form solutions of exchange and interest rate dynamics prior to the entrance into a currency union. These two-country stochastic equilibrium models will form the theoretical basis of this paper.

The European Monetary Union (EMU) started on 1 January 1999 with a first wave of 11 countries, namely Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain. Exactly two years later, Greece joined the EMU and it is precisely this case that we will focus on in our empirical analysis.

Obviously, the economic relevance of testing the aforementioned theoretical exchange and interest rate models empirically on the basis of historical data seems reasonable in view of at least two political prospects in the near future: (a) the enlargement of the group of current EMU members, and (b), multilateral efforts to create currency unions in other parts of the world (e.g. in the Mercosur area or in East Asia).[1]

The potential gains of collecting insights from EMU experiences for such future events are at least twofold: (a) For the conduct of an optimal monetary, exchange and interest rate policy during the transition into the currency union. And (b), for calibrating the theoretical exchange and interest rate models with the aim of applying them in a broad range of financial areas, e.g. in the selection of optimal hedging strategies for risk-averse investors or in the valuation of exchange- and interest-rate sensitive claims.

The paper is organized as follows: Section 2 reviews the aforementioned theoretical models on exchange and interest rate dynamics. These models have several empirical implications which should be observable in real-world data. In Section 3 the empirical implications are tested on the basis of Greek exchange and interest rate data vis-a-vis (initial) Euroland for the time prior to Greek's EMU-entrance on 1 January 2001. It is shown that the theoretical models can explain many empirical features of the Greek variables. Section 4 offers some concluding remarks.

[1] See, among others, Eichengreen (1999, p. 95).

2 Exchange and interest rate dynamics

The aim of this section is twofold. First, we present a stochastic equilibrium model of exchange rate dynamics prior to a monetary union within a two-country framework. Second, we derive a model of interest rate convergence based on the corresponding exchange rate dynamics under the additional assumption that a special form of the uncovered interest parity may hold between the two countries prior to the currency union.

2.1 Exchange rate dynamics

For notational convenience, let the political and/or monetary authorities of the two countries involved officially announce at date t_A their aim of entering a bilateral currency union from the future date t_S onwards ($t_S > t_A$). Apart from the specific timing, let the authorities also announce at date t_A the final conversion rate $\overline{x}$ between the two currencies at which both countries will enter the currency union.

As a general law of exchange rate valuation, valid under each exchange rate system as well as under each sequence of exchange rate regimes, consider the well-known stochastic version of the (continuous-time) monetary exchange rate model with flexible prices. In this model, the logarithmic exchange rate at time t, $x(t)$, equals the sum of a macroeconomic fundamental $k(t)$ plus a speculative term proportional to the expected rate of change in the exchange rate:[2]

$$x(t) = k(t) + \alpha \cdot \frac{E[dx(t)|\phi(t)]}{dt}, \quad \alpha > 0. \tag{1}$$

In Eq. (1), $E[\cdot|\cdot]$ denotes the usual expectation operator conditional on the present time-t information set $\phi(t)$ which includes all information available to rational market participants at time t. To simplify, the composite fundamental k, which consists of several economic variables such as domestic and foreign money supplies and outputs, can be thought of as a collection of all economic and/or political factors which market participants deem to be important for the market valuation of exchange rates.

On the analogy of EMU, assume that both economies maintain a managed-float exchange rate regime before t_S. Such a situation might be considered comparable to the European Exchange Rate Mechanism (ERM) before 1999 or the ERM-II from 1999 onwards for those countries which did not belong to the first wave of EMU-ins.[3] Fol-

[2] For an overview and theoretical appreciation see Bertola (1994).

[3] Strictly speaking the ERM and ERM-II are target zone regimes with well defined central parities and intervention points. Nevertheless, because of the relatively wide bandwidths (±15% around the

lowing the lines of Wilfling (2002) such a managed-float system is adequately modelled by letting the economic fundamental k in Eq. (1) follow an Ornstein-Uhlenbeck process with stochastic differential

$$dk(t) = \eta \cdot [\overline{x} - k(t)] \cdot dt + \sigma \cdot dw(t), \tag{2}$$

where $\sigma > 0$ denotes the infinitesimal standard deviation and $dw(t)$ the increment of a Wiener process. The quantity $\eta \cdot [\overline{x} - k(t)]$ represents the force that keeps pulling the fundamental towards its long-run target value $\overline{x}$ with magnitude proportional to the current deviation of the process from the conversion rate $\overline{x}$. The parameter $\eta \geq 0$ indicates the strength of the target-reversion property. Therefore, η can be interpreted as a measure of the willingness and/or the capability of the central banks to stabilize the exchange rate x near the target parity by appropriate interventions in foreign exchange markets.[4]

On modelling the stochastic dynamics of the macroeconomic fundamental k by the Ornstein-Uhlenbeck process (2), the exchange rate equation (1) represents a stochastic differential equation (s.d.e.). Technically, this s.d.e. can be solved by Ito-integration methods and the imposition of adequate constraints correctly reflecting the economic requirements at hand.

In a first step, consider the time interval $[0, t_A)$ so that the political aim of entering the currency union has not yet been officially announced by the authorities. For simplicity, assume for a moment that agents therefore expect the present managed-float regime to hold for the indefinite future.[5] Following the lines of Wilfling (2002), the (bubble-free) solution of the s.d.e. (1) can then be obtained as

$$x(t) = \overline{x} + \frac{k(t) - \overline{x}}{1 + \alpha\eta} = \frac{1}{1 + \alpha\eta} \cdot k(t) + \frac{\alpha\eta}{1 + \alpha\eta} \cdot \overline{x}. \tag{3}$$

Obviously, under a managed-float which is considered as permanent by market par-

central parities) and the widespread tendencies of the central banks to intervene intramarginally, such target zone systems may well be approximated by managed-float systems (for a justification see Svensson, 1992, p. 134).

[4] As two special cases, consider the parameter choices $\eta = 0$ and $\eta \to \infty$. For $\eta = 0$, central banks refrain from intervening in foreign exchange markets. In this case, the Ornstein-Uhlenbeck process in Eq. (2) becomes a Brownian motion and the exchange rate system prior to the currency union is identified as a free-float. For $\eta \to \infty$, any exchange rate deviation from the conversion rate $\overline{x}$ will instantaneously be corrected for by appropriate interventions. Hence, the two economies *de facto* live in a fixed-rate system prior to the currency union.

[5] Evidently, agents in foreign exchange markets are forward-looking and rational so that they will form anticipations about the prospective currency union before the official political announcement. In the empirical analysis below, the rather unrealistic assumption that the announcement at date t_A provides complete 'news' for market participants will be relaxed.

ticipants, the exchange rate x is a weighted average of the fundamental k and the conversion rate $\overline{x}$.

Next, consider the interim period $[t_A, t_S)$. The aim of entering a currency union at t_S on the basis of the conversion rate $\overline{x}$ has now been officially announced. Nevertheless, market participants may deem the political announcement more or less credible, in that they consider a delay in (or, even more extreme, a complete surrender of) the currency union possible. Along the same lines as Wilfling and Maennig (2001) such a lack of credibility can be modelled by assuming that market participants – based on their present date-t information set $\phi(t)$ – associate a specific parametric probability distribution function with the lifetime of the current managed-float exchange rate system. Denoting this system-lifetime by Z, the probability that Z does not exceed the future date $s > t$ can reasonably be modelled by

$$F_Z(s; p, \lambda) \equiv \Pr\{Z \leq s | \phi(t)\} = \begin{cases} 0 & \text{for} \quad s < t_S \\ 1 - p \cdot \mathrm{e}^{\lambda(t_S - s)} & \text{for} \quad s \geq t_S, p \in [0,1], \lambda \geq 0 \end{cases} . \quad (4)$$

The parameters p and λ in the distribution function (4) have neat economic interpretations which stem from reliability theory of technical systems. These properties are described in detail in Wilfling and Maennig (2001) and Wilfling (2002, 2003b). For the purpose of this paper it is important to note two special cases of the distribution function (4) as well as another important characteristic of this credibility modelling:

(a) For $p = 0$ (λ arbitrarily chosen) the fixing date t_S is considered fully credible by the market. In this case agents expect the entrance into the currency union to be executed punctually at date t_S.

(b) For $(p, \lambda) = (1, 0)$ agents believe that the entrance into the union will never take place, i.e. the current managed-float system will be permanent forever.

(c) It should be noted that agents are free to revise their credibility assessment at any point in time. These revisions may be due to political and/or economic news arriving on the scene. Any modified assessment will be reflected by corresponding variations in the parameters p and λ. The possibility of continuously revising the uncertainty parameters p and λ will be important in the subsequent empirical analysis.

On the basis of the uncertainty structure (4) Wilfling (2002) derives the equilibrium exchange rate path for $t \in [t_A, t_S)$ as

$$x(t) = \overline{x} + \frac{k(t) - \overline{x}}{1 + \alpha\eta} \cdot \left[1 - \left(1 - \frac{p \cdot (1 + \alpha\eta)}{1 + \alpha\eta + \alpha\lambda} \right) \cdot \mathrm{e}^{(1+\alpha\eta)\cdot(t - t_S)/\alpha} \right] . \quad (5)$$

Finally, let us assume that the monetary authorities — in spite of potential temporary scepticism—in fact implement the currency union at date t_S. Under this circumstance, the exchange rate for $t \geq t_S$ will be given by

$$x(t) = \overline{x}. \tag{6}$$

2.2 Interest rate dynamics

The exchange rate dynamics presented in the previous section now allows us to derive a complete term structure of interest rate differentials for the two economies involved. For this, assume that all conditions necessary to pledge the uncovered interest parity between the two economies to hold at all points in time are satisfied.[6]

Denoting domestic and foreign interest rates by i and i^*, respectively, and the term of any interest rate by τ, an approximate form of the uncovered interest parity condition at date t is given by

$$ID(t,\tau) \equiv i(t,\tau) - i^*(t,\tau) = \frac{E[x(t+\tau)|\phi(t)] - x(t)}{\tau}. \tag{7}$$

Obviously, in equilibrium the differential between domestic and foreign interest rates of the same term τ must equal the expected rate of currency depreciation accrued over the future time interval $(t, t+\tau]$.

Now, using the exchange rate path from the equations (3), (5) and (6), the corresponding interest rate differentials are derived using basic principles of stochastic calculus. For example, the interest rate differential for $t < t_A$ under a permanent managed-float (3) is easily computed as

$$ID(t,\tau) = \frac{\overline{x} - k(t)}{1+\alpha\eta} \cdot \frac{1 - \mathrm{e}^{-\eta\cdot\tau}}{\tau}.^{7} \tag{8}$$

Next, consider the differential dynamics during the interim period $[t_A, t_S)$. Again,

[6] In particular, international investors consider home and foreign bonds (of the same term) as perfect substitutes. Moreover, both countries are linked by perfect international capital mobility. A further assumption, implicit in the theoretical interest rate models below, is that the domestic economy is assumed to be small. Thus the domestic economy is not able to influence the foreign interest rates by economic policy but has to accept the foreign rates as exogenously given (cf. Wilfling, 2003a; 2003b).

[7] See Wilfling (2003a, formula 15). In the same paper Wilfling also considers so-called instantaneous interest rate differentials, i.e. differentials for extremely short-term interest rates ($\tau \to 0$). Instantaneous interest rates (which may be thought of as overnight rates) exhibit a number of singularities as compared with interest rates of strictly positive terms. Instantaneous differentials will be discussed in Section 3.2.

potential market scepticism about the punctual implementation of the currency union must be taken into account by incorporating the uncertainty structure (4) into the computation of all interest rate differentials. However, for reasons which will become evident below, assume that the political announcement of entering the union at date t_S is considered fully credible by market participants.[8] Technically, this market assessment is reflected by setting $p = 0$ in Eq. (4). Furthermore, making use of the exchange rate paths (5) and (6) a two-branched interest rate differential path obtains as

$$ID_1(t,\tau) = \left[\frac{k(t)-\overline{x}}{1+\alpha\eta}\right] \cdot \left\{\frac{e^{-\eta\cdot\tau}-1}{\tau} + \left[\frac{1-e^{\tau/\alpha}}{\tau}\right] \cdot e^{(1+\alpha\eta)\cdot(t-t_S)/\alpha}\right\} \tag{9}$$

for $t \in [t_A, t_S - \tau)$, while for $t \in [t_S - \tau, t_S)$ the differential is given by

$$ID_2(t,\tau) = \frac{1}{\tau} \cdot \left[\frac{\overline{x}-k(t)}{1+\alpha\eta}\right] \cdot \left[1 - e^{(1+\alpha\eta)\cdot(t-t_S)/\alpha}\right]. \tag{10}$$

Finally, consider the interest rate dynamics for the time after t_S. Both currencies have been irreversibly fixed at the conversion rate $\overline{x}$. Consequently, the expected rate of change in the exchange rate will be zero over any future time interval. Hence, the uncovered interest parity condition (7) implies zero-differentials for arbitrary terms, i.e. for $t \geq t_S$ we have

$$ID(t,\tau) = 0 \tag{11}$$

for arbitrary term $\tau > 0$.

3 Empirical results

In this section, we analyze Greek exchange and interest rate data vis-a-vis the corresponding Euroland variables. For this we proceed in two steps. (a) We further explore the dynamic equations from the previous section in order to derive implications which should be observable empirically. (b) We verify these implications empirically thereby indirectly confirming the relevance of the theoretical dynamic models.

3.1 Exchange rates

Although the exchange rate path sequences (3), (5) and (6) offer several implications which could be analyzed empirically, we restrict attention to the most important con-

[8] Closed-form solutions for interest rate differentials of arbitrary term under a lack of perfect credibility can be derived by analogous techniques as in Wilfling (2003b). Explicit formulae are available upon request.

sequence: the process of (conditional) variances of exchange rate returns before and during the interim period $[t_A, t_S)$. To motivate, consider a useful volatility measure of the (logarithmic) exchange rate x: the infinitesimal variance (subsequently denoted by $\nu^2_{\{x\}}$) which can be computed by the well-known *Ito*-lemma. Wilfling (2002) obtains the infinitesimal variances of the exchange rate paths (3) and (5) as

$$\nu^2_{\{x\}} = \left[\frac{\sigma}{1+\alpha\eta}\right]^2 \quad \text{for } t < t_A, \tag{12}$$

and

$$\nu^2_{\{x\}} = \left[\frac{\sigma}{1+\alpha\eta}\right]^2 \cdot \left[1 - \left(1 - \frac{p\cdot(1+\alpha\eta)}{1+\alpha\eta+\alpha\lambda}\right)\cdot \mathrm{e}^{(1+\alpha\eta)\cdot(t-t_S)/\alpha}\right]^2 \quad \text{for } t \in [t_A, t_S), \tag{13}$$

while the infinitesimal variances are constantly equal to zero for $t \geq t_S$.

Note that the exchange rate variances are constant before t_A while they are time-dependent (non-stationary) during the interim period $[t_A, t_S)$. Furhtermore, the uncertainty parameters p and λ directly enter the volatility path (13). It is easy to verify the following two results by means of standard calculus:

(a) Apart from the parameter constellation $(p, \lambda) = (1, 0)$, the interim volatility path (13) always lies below the variance path (12).

(b) The infinitesimal variances along the interim path (13) are, *ceteris paribus*, increased by increases in the uncertainty parameter p and/or reductions in λ while the variances decrease for reductions in p and/or increases in λ.[9]

Figure 1 depicts the qualitative nature of the complete infinitesimal variance path $\nu^2_{\{x\}}$ for the given set of structural parameters $\alpha = 1.0, \eta = 0.5, \sigma = 2.0, t_A = 1.0$, $t_S = 2.0$. Higher market uncertainty about the punctual implementation of the currency union cleary leads to an outward shift of the infinitesimal variances as opposed to the certainty case represented by $p = 0$ (λ arbitrarily chosen). In the most extreme setting market participants believe that the union will never be implemented and set $(p, \lambda) = (1, 0)$. In this case, the volatility path during the interim period $[t_A, t_S)$ is

[9] The relevance of this last point is motivated by noting that the lifetime distribution (4) characterizing the credibility of the political announcement of future entrance into the currency union—has expected value and variance given by $E[Z|\phi(t)] = t_S + p/\lambda$ and $\mathrm{Var}[Z|\phi(t)] = p\cdot(2-p)/\lambda^2$, respectively. It is easy to check that both quantities, *ceteris paribus*, are strictly increasing in p and strictly decreasing in λ. In other words, raising p and/or lowering λ pushes the expected date of implementing the currency union into the more distant future and induces markets to attach a higher variance to the entering-date. This loss of credibility unambiguously entails higher exchange rate variances during the interim period.

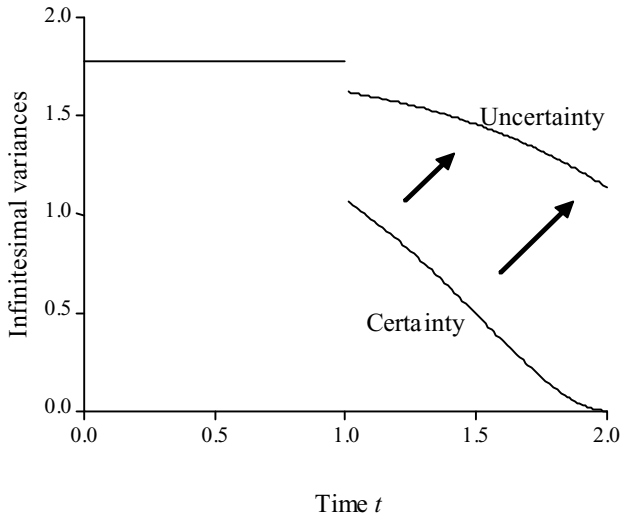

Figure 1: Infinitesimal variance-paths of the logarithmic exchange rate

simply the continuation of the horizontal variance line on the interval $[0, t_A)$ as can be verified formally from the volatility paths (12) and (13). However, it must be emphasized once more that due to news upcoming during the interim period, agents may revise their credibility assessments and thus switch to different p- and λ-values. In this case the volatility path between t_A and t_S is not unique in the sense that—in reconciliation with the switches to modified p- and λ-values infinitesimal variances jump between alternative volatility curves.

These theoretical considerations imply a clear-cut volatility pattern for the exchange rate x before and during the interim period. The infinitesimal variances should switch between two successive volatility regimes: on the interval $[0, t_A)$ variances should be constantly high followed by a regime of lower and declining variances during the interim period $[t_A, t_S)$.

For mathematical reasons explained in Wilfling (2002) the infinitesimal variance of the logarithmic spot rate x at time t given by the volatility paths (12) or (13) is a good proxy for (and is well approximated by) the conditional variance of one-step-ahead exchange rate returns, i.e.

$$\nu^2_{\{x\}}(t) \approx \mathrm{Var}\left[x(t+1) - x(t)|\phi(t)\right]. \tag{14}$$

It can be shown that the conditional variances of the exchange rate returns must reveal the same qualitative features as the infinitesimal variances. Consequently, the conditional variances of the one-step-ahead exchange rate returns should undergo exactly

the same shift in volatility regime as elaborated above for the infinitesimal variances $\nu^2_{\{x\}}$.[10]

Before running the empirical analysis one further issue must be addressed. The theoretical exchange rate model from above implicitly assumes the existence of a clear-cut announcement date t_A from which onwards rational market participants fully incorporate their knowledge of the prospective currency union into currency valuation. Furthermore, the date t_A is assumed to have news-character, i.e. before t_A agents are not aware that an announcement will be made and consequently are completely surprised by the political announcement.

Clearly, in reality agents anticipate the prospective currency union long before any official announcement and incorporate this knowledge immediately into their valuation schemes. Hence, the question to be answered by our empirical analysis below will be: When did foreign exchange markets begin to calculate with the Greek EMU-entrance? Econometrically, we will answer this question by analyzing the volatility structure of the exchange rate returns of the Greek drachma *vis-a-vis* the Euro with the hope of finding a statistically significant shift from a high to a low-volatility regime at that moment when markets incorporate the prospective EMU-entrance into currency pricing. Following this line of argument, the main purpose of our empirical exchange rate analysis is to identify the beginning of the interim period between the successive exchange rate regimes.

The exchange rate data used in this study are daily spot rates of the Greek Drachma (GRD) *vis-a-vis* the euro covering the period from 15 December 1998 until 31 December 2000. The rates—measured as GRD-prices of 1 euro—were collected from the historical database provided by the *OANDA-FXTrade*-website (*http://www.oanda.com/convert/-fxhistory*) and are daily averages of interbank rates recorded at seven days per week.[11]

Figure 2 depicts the nominal spot rates during the sampling period along with the daily exchange rate returns R_t defined as

$$R_t = 100 \cdot (x_t - x_{t-1}) \tag{15}$$

(first differences in logarithmic rates). The graph of the nominal rates also contains the

[10] A mathematical description of the approximation accuracy is given in Wilfling (2002, Eqs. 9 to 12) or in Karlin and Taylor (1981, p. 159).

[11] Note that the first rate was recorded on 15 December 1998, i.e. two weeks before the official start of Stage III of EMU. This is the first GRD-euro exchange rate provided by the *OANDA* currency converter. The dataset contains two evident outliers for the dates 24/09/2000 and 25/09/2000 (identified by visual inspection). These observations were excluded from the analysis. After exclusion, the dataset consists of 746 observations.

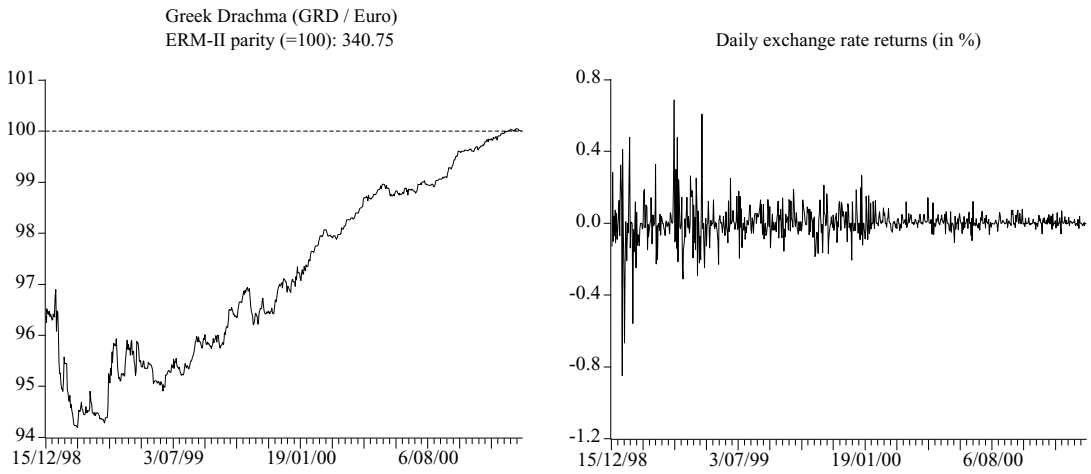

Figure 2: Nominal drachma-euro spot rates and daily exchange rate returns

ERM-II central parity (the later conversion rate) which was scaled to 100. Obviously, the exchange rates converge towards the conversion rate at the end of the interim period. It is interesting to note that this convergence may be explained in at least two different ways:

(a) Consider the exchange rate equation (5) and assume that agents become more and more convinced about the punctual implementation of the monetary union during the interim period $[t_A, t_S)$. Consequently, they set $p = 0$ at some date before t_S and retain this assessment until t_S. It is evident from Eq. (5) that in this case the exchange rate x converges (with probability 1) towards $\overline{x}$ for $t \to t_S$. Here the convergence towards the conversion rate is due to agents' removal of arbitrage opportunities during the transition into the currency union.

(b) On setting $p = 0$ in Eq. (5), it is easy to check that the exchange rate x can be pushed arbitrarily close towards $\overline{x}$ at any date $t \in [t_A, t_S)$ by choosing a sufficiently high value of η, or in other words, by an appropriately high degree of central bank intervention.

In the end, an exact *ex-post* determination of the cause of exchange rate convergence has to be explored on the basis of intervention data of both central banks involved.

The daily exchange rate returns reveal a heteroscedastic pattern consistent with the theoretical exchange rate model from above. At the beginning of the sampling period the returns exhibit high volatility, switching to lower variances at that moment (or during that time window) from which onwards market participants definitely incorpo-

rate the prospective currency union into exchange rate valuation. The date (or the short time interval) on which this switch happens can be interpreted as the empirical equivalent to the announcement date t_A from the theoretical model.

The appropriate econometric technique for analyzing volatility shifts of the required type are provided by Markov-switching (or regime switching) models. In spite of some early methodological contributions to Markov-switching models their modern formulation is due to Hamilton (1988, 1989). Hamilton and Raj (2002) provide an extensive overview of Markov-switching applications in economics and finance. In our analysis we make use of the switching model elaborated in Wilfling (2002) which primarily constitutes a Markov-switching GARCH model as developed in Gray (1996b) but adapted for t-distributed exchange rate returns within each regime.[12]

To set up the model, recall first the probability density function of a displaced t-distribution with ν degrees of freedom, mean μ and variance h:

$$t_{\nu,\mu,h}(x) = \frac{\Gamma[(\nu+1)/2]}{\Gamma[\nu/2] \cdot \sqrt{\pi \cdot (\nu-2) \cdot h}} \cdot \left[1 + \frac{(x-\mu)^2}{h \cdot (\nu-2)}\right]^{-(\nu+1)/2} \tag{16}$$

with $\Gamma(z) \equiv \int_0^\infty l^{z-1} \cdot \mathrm{e}^{-l}\, dl, z > 0$, denoting the complete gamma function.

The general idea of a univariate Markov-switching model is that the data generating process may be affected by a non-observable random variable S_t representing the state the data generating process is in at date t. In our analysis the state variable $S_t, t = 0, 1, \ldots, T$, can take on two distinct values: $S_t = 1$ indicates that the data generating process is in the high volatility regime whereas for $S_t = 2$ the generating process is in the low volatility regime. Translated into our exchange rate framework, Regime 1 characterizes the situation in which market participants have not yet incorporated the currency union into their valuation scheme, whereas in Regime 2 they anticipate the future exchange rate fixing at the conversion rate implying that the returns must exhibit low volatility.

The complete Markov-switching model is now set up by specifying parametric forms for the conditional means and variances which the return R_t may take on conditional upon the regime indicator $S_t = i, i = 1, 2$. Denoting the mean and variance within regime i by μ_{it} and h_{it}, respectively, the conditional distribution of the return R_t can

[12]The use of t-distributed rather than normally distributed exchange rate returns within each regimes may be motivated by the 'fat-tail'-property of many financial variables such as stock price changes or exchange rate returns (see Bollerslev, 1987).

be represented as a mixture of two displaced t-distributions:

$$R_t|\phi_{t-1} \sim \begin{cases} t_{\nu_1,\mu_{1t},h_{1t}} & \text{with probability } p_{1t} \\ t_{\nu_2,\mu_{2t},h_{2t}} & \text{with probability } (1-p_{1t}) \end{cases}, \tag{17}$$

where $p_{1t} \equiv \Pr\{S_t = 1|\phi_{t-1}\}$ denotes the so-called *ex-ante* probability of being in Regime 1 at date t.

As in Wilfling (2002) we model the regime-specific mean μ_{it} as a parsimonious first-order autoregressive process (AR(1) process), i.e.

$$\mu_{it} = a_{0i} + a_{1i} \cdot R_t \qquad \text{for } i = 1, 2. \tag{18}$$

In contrast to the conditional mean μ_{it}, finding an adequate functional form for the regime-specific variance h_{it} turns out to be more problematic. The reason for this complication—known as a phenomenon called path dependence—is explained in Gray (1996b). Without going into further technical details we simply adopt Wilfling's (2002) approach which itself is built on Gray's (1996b) general solution of the path-dependence problem in regime switching models—and observe that from Eq. (17) the variance of the return R_t can be written as

$$\begin{aligned} h_t &= E\left[R_t^2|\phi_{t-1}\right] - \{E\left[R_t|\phi_{t-1}\right]\}^2 \\ &= p_{1t} \cdot \left(\mu_{1t}^2 + h_{1t}\right) + (1-p_{1t}) \cdot \left(\mu_{2t}^2 + h_{2t}\right) - \left[p_{1t} \cdot \mu_{1t} + (1-p_{1t}) \cdot \mu_{2t}\right]^2. \end{aligned} \tag{19}$$

The quantity h_t can be thought of as an aggregate of conditional variances from both regimes and now provides the basis for the specification of the regime-specific conditional variances $h_{it+1}, i = 1, 2$, in the form of parsimonious GARCH(1,1) models. To be more explicit, consider the following GARCH(1,1) variance process within regime i:

$$h_{it} = b_{0i} + b_{1i} \cdot \epsilon_{t-1}^2 + b_{2i} \cdot h_{t-1}, \tag{20}$$

with h_{t-1} given by Eq. (19) lagged one period, while ϵ_{t-1} is obtained from

$$\begin{aligned} \epsilon_{t-1} &= R_{t-1} - E\left[R_{t-1}|\phi_{t-2}\right] \\ &= R_{t-1} - \left[p_{1t-1} \cdot \mu_{1t-1} + (1-p_{1t-1}) \cdot \mu_{2t-1}\right]. \end{aligned} \tag{21}$$

To close the model, it remains to specify the transition probabilities of the regime indicator S_t. For simplicity, consider a first order Markov process with the constants

$\pi_1, \pi_2 \in [0,1]$ and define

$$\begin{aligned}
&\Pr\{S_t = 1 | S_{t-1} = 1\} = \pi_1, \\
&\Pr\{S_t = 2 | S_{t-1} = 1\} = 1 - \pi_1, \\
&\Pr\{S_t = 2 | S_{t-1} = 2\} = \pi_2, \\
&\Pr\{S_t = 1 | S_{t-1} = 2\} = 1 - \pi_2.
\end{aligned} \tag{22}$$

Replicating the derivations in Gray (1996b, p. 58) it is now possible to derive the log-likelihood function of the Markov-switching GARCH(1,1) model as specified in the Eqs. (16) to (22). Formally, the likelihood function contains the *ex-ante* probabilities $p_{1t} = \Pr\{S_t = 1 | \phi_{t-1}\}$. The whole series of *ex-ante* probabilities can be estimated recursively using the t-densities from Eq. (16). Exact formulas for the log-likelihood function and the recursive estimation procedure for the p_{1t}'s are given in Wilfling (2002). Here, we simply list all parameters in regime i ($i = 1, 2$), which have to be estimated according to the specifications (16) to (22): the AR(1) mean parameters a_{0i}, a_{1i}, the GARCH(1,1) variance parameters b_{0i}, b_{1i}, b_{2i}, the degrees of freedom from the t-distribution within regime i, ν_i, and the transition probabilities π_i.

Table 1 in the appendix presents the maximum-likelihood estimates of the Markov-switching model from the Eqs. (16) to (22) for the returns of the Greek drachma *vis-a-vis* the euro. Maximization of the log-likelihood function was performed by the 'MAXIMIZE'-routine within the software package RATS 5.02 using the BFGS-algorithm, heteroscedasticity-consistent estimates of standard errors and suitably chosen starting values for all parameters involved.

The estimates in Table 1 can be analyzed and interpreted economically, but we waive to go into details here and only briefly mention three aspects:

(a) The estimates of the degrees-of-freedom parameters ν_1 and ν_2 are both larger than 2.0, an important stipulation for the existence of the variance of a t-distribution. However, since both ν-parameters are smaller than 4.0, the estimated t-distributions within each regime have infinite fourth moments hinting at excess kurtosis for the exchange rates returns of the Greek drachma.

(b) The constant transition probabilities π_1 and π_2 are very close to one. Since both quantities represent the probability of the data generating process staying in the same regime during the transition from date $t-1$ to t (no structural break between $t-1$ and t), both volatility regimes reveal a high degree of persistence.

(c) The lower part of Table 1 contains a diagnostic check of the model fit by pro-

viding Ljung-Box statistics for serial correlation of the squared (standardized) residuals out to the lags $1, 2, 3, 5$ and 10. Obviously, the null hypothesis of no autocorrelation cannot be rejected out to all lags at any conventional level. This provides some econometric evidence in favour of our Markov-switching AR(1)-GARCH(1,1) specification.

Finally, let us address two conditional probabilities which are of high inferential relevance for the question of when financial markets began to incorporate the currency union into their pricing schemes. The series of *ex-ante* probabilities $\Pr\{S_t = 1|\phi_{t-1}\}, t = 2, ..., T$, can be estimated recursively as mentioned above, while the series of the so-called *smoothed* probabilities, $\Pr\{S_t = 1|\phi_T\}, t = 1, ..., T$, have to be computed after model estimation by the use of filter techniques.[13]

The *ex-ante* probabilities are useful in forecasting the one-step-ahead regimes based on an information set which evolves over time. Due to their definition, the *ex-ante* probabilities reflect current market perceptions of the one-step-ahead volatility regime. Consequently, these quantities represent a good measure of market sentiments with respect to the question if and to what extent the prospective exchange rate fixing at the conversion level will be incorporated into currency pricing in the next step. In contrast to the *ex-ante* probabilities, the *smoothed* probabilities provide a basis for inferring *ex post* if and when regime switches have occurred in the sample.

Figure 3 displays the *ex-ante* as well as the *smoothed* probabilities in one graph. The lower part of Figure 3 also depicts the estimated conditional variance process over the whole sample. It is most interesting to interpret the evolution of both regime-1 probabilities. According to the theoretical exchange rate model from Section 3.1, both series of regime-1 probabilities should reveal the following clear-cut pattern over time: Under a fully credible announcement at t_A of entering the currency union from t_S onwards, the probabilities should take on the value 1 (representing the high-volatility regime of exchange rate returns) and should then immediately drop to zero at t_A from which onwards financial markets begin to incorporate the prospective union into their pricing schemes (i.e. during the low-volatility regime). However, the theoretical model also provides an explanation for temporary deviations from the strict evolution along to the one- and zero-baselines. If market participants deem a delay in the exchange rate regime switch from floating to fixed rates possible, they reflect this lack of credibility by assigning appropriate values to the uncertainty parameters p and λ in the equilibrium interim exchange rate path (5). According to the explanations in Section

[13]The *smoothed* probabilities for the Greek drachma were computed on the basis of an filter algorithm provided by Gray (1996a).

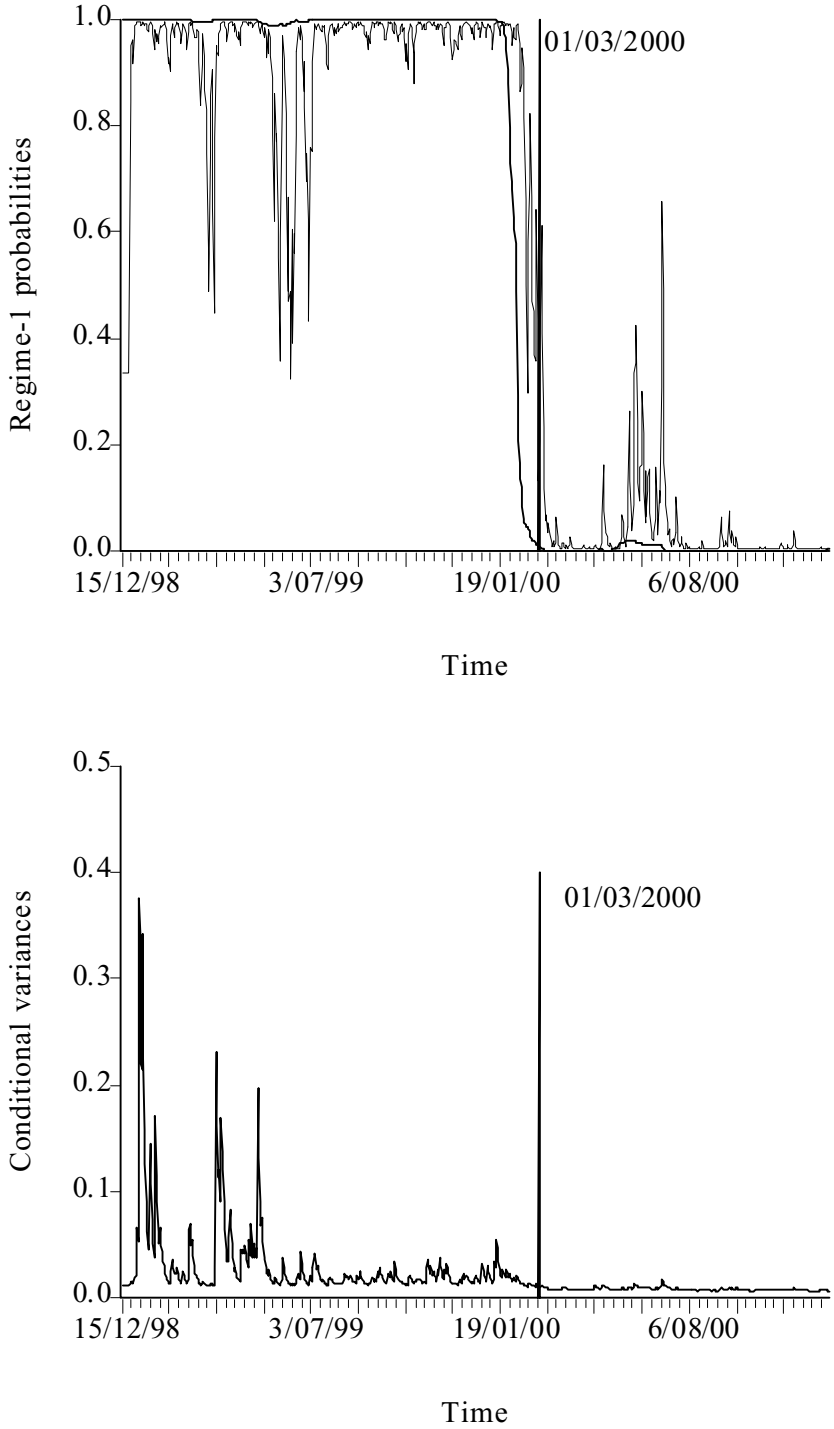

Figure 3: *Ex-ante* (thin line), smoothed (bold line) probabilities and conditional variances of GRD-returns

3.1, these variations in the uncertainty parameters may temporarily reduce or increase the variances of exchange rate returns thus giving room for statistically significant switches between the alternative volatility regimes (see Figure 1). Such a scenario of political uncertainty is very likely to disturb the clear-cut pattern in the evolution of the regime-1 probabilities by up- and downturns from the one- and zero-baselines.

As expected, the *ex-ante* probabilities exhibit a more volatile pattern than the corresponding *smoothed* probabilities. The reason for these sharper deviations from the one- and zero-baselines evidently lies in the more restrictive information set ϕ_{t-1} used to compute the *ex-ante* probabilities. Nevertheless, both regime-1 probabilities behave exactly as predicited by the theoretical model. According to this analysis, the low-volatility regime is unambiguously entered around 1 March 2000 (see the marker in Figure 3). The conditional variances in the lower part of Figure 1 confirm this empirical finding. Therefore, in the subsequent econometric analysis of interest rates we will treat the period between 1 March 2000 and 1 January 2001 as the empirical equivalent to the theoretical interim period $[t_A, t_S)$.

3.2 The term structure of interest rate differentials

Having detected the interim period between 1 March 2000 and 1 January 2001, we will now analyse the interest rates of both economies. Figure 4 displays 1-day (overnight) as well as 1-, 2- and 3-months interest rates for Greece and Euroland together with the corresponding differentials during the interim period (219 observations). All rates were collected from *Datastream* (EURIBOR). At the beginning of the interim period, Greek interest rates were much higher than the corresponding rates of Euroland for all maturities. All interest rate differentials start between 5 and 6 percent before they all gradually shrink to zero at the Greek EMU-entrance.

Formally, the interest rate differential convergence towards zero at the end of the interim period for arbitrary term $\tau > 0$ follows from the ID_2 branch in Eq. (10) for which we have (with probability 1)

$$\lim_{t \to t_S} ID_2(t, \tau) = 0.$$

Economically, the equalization of interest rates of the same term can be explained by an arbitrage argument: if interest rates of the same term did not equalize completely at the moment of transition into the currency union, riskless profit opportunities would exist infinitesimally shortly before t_S by buying Greek bonds and selling Euroland-bonds (or vice versa). The only way to rule out these arbitrage transactions is a complete

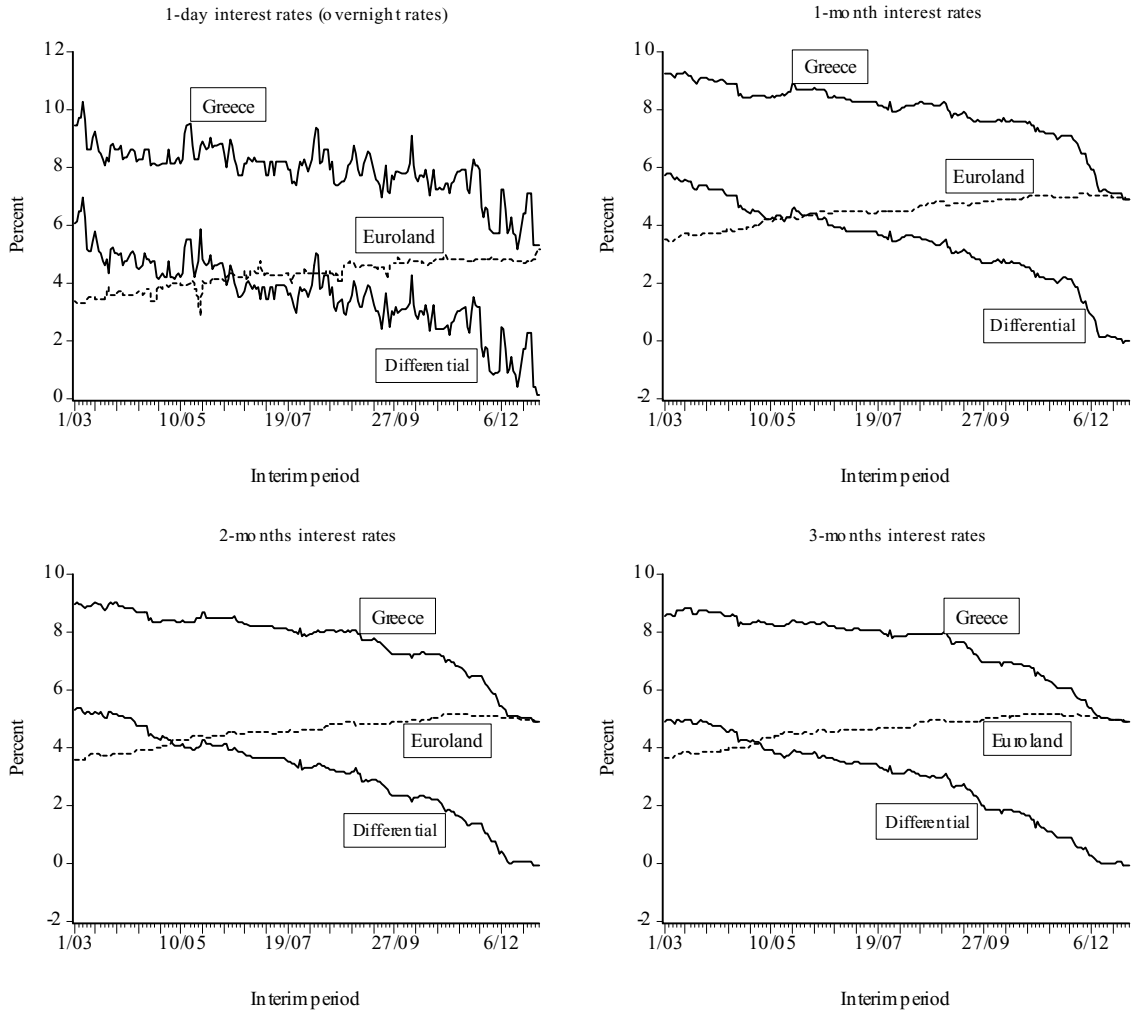

Figure 4: Interest rates and differentials of different maturities

equalization of interest rates.

Observe, however, that the overnight interest rate differential — in contrast to all other differentials — appears more volatile and only reaches the zero level shortly before 1 January 2001. This late convergence of extremely short-term differentials can be given an appealing explanation by looking at the theoretical instantaneous interest rate differential (see Footnote 7). The instantaneous differential $ID(t,0), t \in [t_A, t_S)$, results from letting $\tau \to 0$ in the ID_1 path (9):

$$ID(t,0) = \lim_{\tau \to 0} ID_1(t,\tau) = \left[\frac{\overline{x} - k(t)}{1+\alpha\eta}\right] \cdot \left[\eta + \frac{1}{\alpha} \cdot \mathrm{e}^{(1+\alpha\eta)\cdot(t-t_S)/\alpha}\right]. \qquad (23)$$

Wilfling (2003a) elaborates some singularities of instantaneous interest rate differen-

tials as opposed to differentials with strictly positive term. An important difference is revealed by evaluating the instantaneous differential at the moment of transition into the currency union:

$$\lim_{t \searrow t_S} ID(t,0) = \left[\frac{\overline{x} - \lim_{t \to t_S} k(t)}{1+\alpha\eta}\right] \cdot \left[\eta + \frac{1}{\alpha}\right].$$

This latter term is different from zero with probability 1, and hence, instantaneous differentials — in contrast to their strictly positive-term counterparts — do (almost surely) not converge to zero at the end of the interim period.

The economic reasoning for this singularity again stems from an arbitrage argument. For any date $t \in [t_A, t_S)$ the maturing date $t+dt$ necessarily falls in the interim period so that the arbitrage argument from above — valid for differentials with strictly positive terms—does not hold anymore. Hence, even for dates shortly before t_S, instantaneous differentials may deviate significantly from zero. For the overnight rates shown in the first panel of Figure 4, the term is strictly positve ($\tau = 1$ day), but — compared with the other terms — rather short. In this sense the overnight differential is similar to the theoretical instantaneous differential attaining the value zero only shortly before 1 January 2001.

In a final step we now address the volatility of interest rate differentials during the interim period. To measure volatility we again invoke the infinitesimal-variance concept as introduced in Section 3.1. Wilfling (2003a) computes the infinitesimal variances of all interest rate differentials during the interim period $[t_A, t_S)$. These time-dependent volatility paths are given by

$$\nu^2_{\{ID(t,0)\}} = \left[\frac{\sigma}{1+\alpha\eta}\right]^2 \cdot \left[\eta + \frac{1}{\alpha} \cdot \mathrm{e}^{(1+\alpha\eta)\cdot(t-t_S)/\alpha}\right]^2 \tag{24}$$

for the instantaneous interest rate differential (23), while for the two differential branches (9) and (10) on the respective domains $[t_A, t_S - \tau)$ and $[t_S - \tau, t_S)$ the infinitesimal variances obtain as

$$\nu^2_{\{ID_1(t,\tau)\}} = \left[\frac{\sigma}{1+\alpha\eta}\right]^2 \cdot \left[\frac{\mathrm{e}^{-\eta\cdot\tau} - 1}{\tau} + \left(\frac{1 - \mathrm{e}^{\tau/\alpha}}{\tau}\right) \cdot \mathrm{e}^{(1+\alpha\eta)\cdot(t-t_S)/\alpha}\right]^2, \tag{25}$$

$$\nu^2_{\{ID_2(t,\tau)\}} = \left[\frac{\sigma}{1+\alpha\eta}\right]^2 \cdot \frac{1}{\tau^2} \cdot \left[1 - \mathrm{e}^{(1+\alpha\eta)\cdot(t-t_S)/\alpha}\right]^2. \tag{26}$$

For a given set of structural parameters, Figure 5 displays the theoretical differential variance paths (24) to (26) for the terms $\tau \to 0$ (instantaneous) and $\tau = 1, 2, 3$ months.

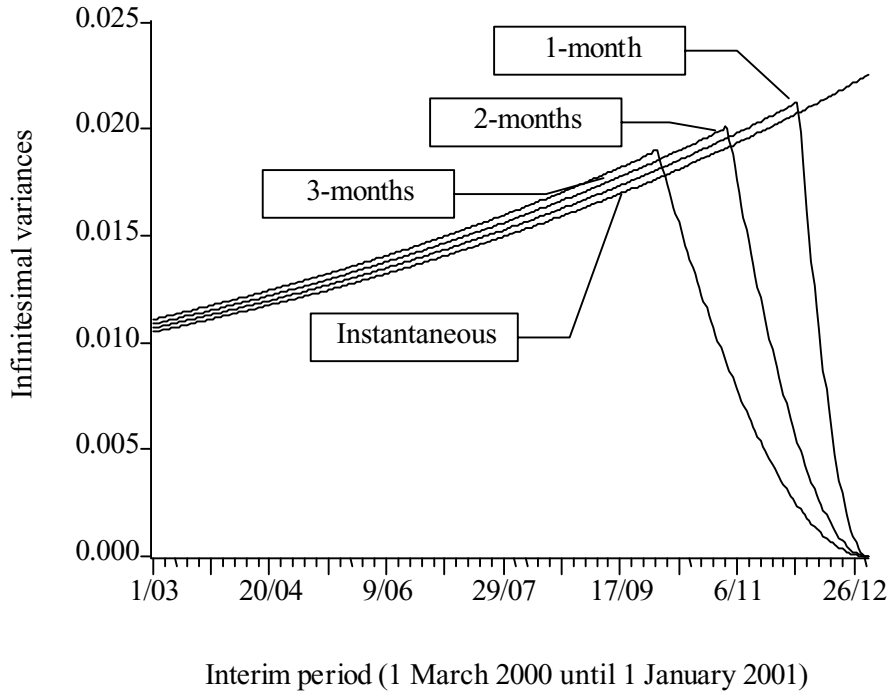

Figure 5: Variance paths of interest rate differentials for different terms

The dynamic properties of these variance paths are described in Wilfling (2003a). It can be shown analytically that the variance path (24) for instantaneous interest rate differentials is monotone increasing during the interim period, while differentials with strictly positive term $\tau > 0$ have a monotone increasing ID_1 variance path (25) on $[t_A, t_S - \tau)$ and a monotone decreasing ID_2 variance path (26) on $[t_S - \tau, t_S)$.

Invoking the same line of argument as in Eq. (14), it follows that the infinitesimal variances from Eqs. (24) to (26) provide accurate proxies for the conditional variances of one-step-ahead changes of the corresponding interest rate differentials, i.e. for all $\tau \geq 0$ we have[14]

$$\nu^2_{\{ID(t,\tau)\}} \approx \text{Var}[ID(t+1,\tau) - ID(t,\tau)|\phi(t)]. \tag{27}$$

Consequently, the qualitative nature of the variance paths (24) to (26) depicted in Figure 5 should be reflected by the conditional variances of the first differences of the empirical interest rate differentials from Figure 4.

[14]For reasons of expositional clarity we do not distinguish notationally between ID_1 and ID_2 differential branches in Eq. (27).

This latter argument suggests the use of suitably specified conventional GARCH models in order to recursively estimate the process of the aforementioned conditional variances.[15] For this, denote the empirical interest rate differential for the term $\tau \geq 0$ at date t by $ID_t(\tau)$ and let the first differences be modelled by

$$\Delta ID_t(\tau) \equiv ID_t(\tau) - ID_{t-1}(\tau) = \boldsymbol{\varphi}_{t-1}^{\top}\mathbf{a} + \epsilon_t. \tag{28}$$

In Eq. (28) $\boldsymbol{\varphi}_{t-1}$ denotes a $(q \times 1)$ vector of explanatory variables whose values are included in the information set ϕ_{t-1} and which may include lagged values of $\Delta ID_t(\tau)$. $\mathbf{a}$ is a $(q \times 1)$ vector of unknown parameters. The disturbance ϵ_t should follow a GARCH(u, v) process, i.e. the distribution of ϵ_t conditional upon ϕ_{t-1} is normal and given by

$$\epsilon_t | \phi_{t-1} \sim N(0, h_t) \tag{29}$$

with

$$h_t = b_0 + \sum_{i=1}^{v} b_i \cdot \epsilon_{t-i}^2 + \sum_{i=1}^{u} c_i \cdot h_{t-i}, \tag{30}$$

where $u, v \geq 0$ represent the order of the GARCH process and the parameters b_i and c_i have to be chosen such that the corresponding variances h_t are positive.

The first practical problem in modelling changes in interest rate differentials is to specify the conditional mean $\boldsymbol{\varphi}_{t-1}^{\top}\mathbf{a}$ in Eq. (28). As in many financial applications we simply use an autoregressive process of order p, i.e.[16]

$$\Delta ID_t(\tau) = a_0 + a_1 \cdot \Delta ID_{t-1}(\tau) + \ldots + a_p \cdot \Delta ID_{t-p}(\tau) + \epsilon_t. \tag{31}$$

The order p was chosen by stepwise downward selection, i.e. starting with a 'high' lag-length p_{max} we step by step reduced the number of regressors until a significant lag was found at the 5% level. After that we fitted a GARCH$(1, 1)$ model, i.e. we reduced the general variance equation (30) to[17]

$$h_t = b_0 + b_1 \cdot \epsilon_{t-1}^2 + c_1 \cdot h_{t-1}. \tag{32}$$

[15] Although GARCH processes are part of the Markov-switching model in Section 3.1, we provide a short description of the conventional GARCH model structure here. For a more general treatment and an early overview of alternative GARCH specifications see Bollerslev et al (1992).

[16] The estimation of AR(p) processes crucially hinges on the stationarity of the time series. Here, we consider first differences of interest rate differentials. For each of these time series the null hypothesis of a unit root is rejected at conventional levels by appropriately specified Augmented-Dickey-Fuller and Phillipps-Perron tests.

[17] For theoretical arguments in favor of a simple GARCH$(1, 1)$ specification, see Bollerslev et al. (1992, p. 10) and the literature cited there.

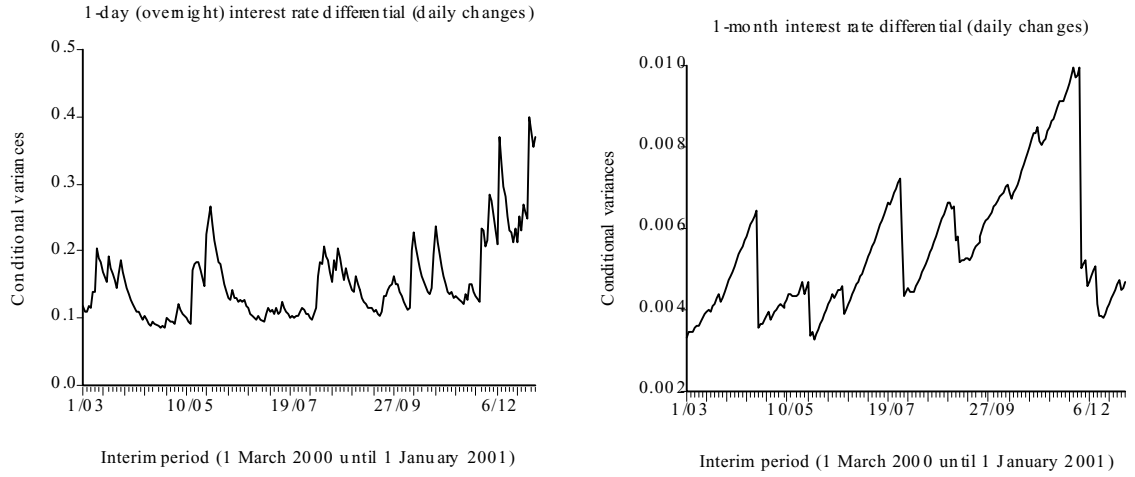

Figure 6: Conditional variances of daily interest rate differential changes

The parameters of the AR(p)-GARCH(1, 1) specifications were estimated by (quasi) maximum likelihood methods using the BHHH-algorithm as implemented in the software package EVIEWS. Heteroscedasticity-consistent standard errors were used to compute z-statistics and p-values (see Bollerslev and Wooldrige, 1992).

Estimation results of the AR(p)-GARCH(1, 1) models for overnight and 1-month interest rate differentials are presented in the appendix of this paper. More informative than the mere parameter estimates are the respective conditional variance processes. These were recursively estimated and are displayed in Figure 6. Evidently, the conditional variances of the overnight differential changes show a clear tendency to increase over time. This qualitative behaviour is largely compatible with the theoretical variance path shown in Figure 5. Also, the conditional variances of the 1-month differential are highly reconcilable with the theoretical volatility path from Figure 5 since these variances exhibit an increasing tendency until 1 December 2000 (representing the ID_1 branch) and after that drop down to a low level (representing the ID_2 branch).

While the empirical results for overnight and 1-month differentials are largely consistent with the implications from the theoretical model, the simple AR(p)-GARCH(1, 1) analysis did not produce similarly reconcilable results for 2- and 3-months interest rate differentials. For these longer terms the estimates are highly sensitive to small modifications in the mean specification. Consequently, we waive to report empirical details here.

4 Concluding remarks

The theoretical models of exchange and interest rate dynamics prior to a currency union presented in Section 2 provide a variety of implications which should be observable in empirical data. The most important implication concerns the pattern of exchange and interest rate volatility-evolution during the interim period. This paper tries to detect these features in exchange and interest rates between Greece and Euroland for the period covering 15 December 1998 and 1 January 2001.

With respect to exchange rates the empirical volatility structure of daily returns is in good accordance with the theoretical model. Based on inferential techniques connected with Markov-switching GARCH models, this circumstance allows us to identify the period between 1 March 2000 and 1 January 2001 as the 'true' interim period, i.e. the time span before the currency union during which financial markets – irrespective of former (in)credible official announcements – incorporated the prospective Greek EMU-entrance into their currency valuation schemes.

During this 'true' interim period the evolution of interest rates only partially corresponds with the theoretical model. First, interest rates of all terms equalize between the two economies and this convergence process takes place exactly in the way predicted by the model with overnight interest rate differentials exhibiting longer lasting volatility than longer-term differentials. Second, the model-inherent volatility structure of interest rate differentials is only reflected by overnight and 1-month interest rates, but not by 2- and 3-months rates.

This latter deficit may be due to various reasons. First, the simple AR(p)-GARCH(1,1) analysis from Section 3.2 may be too simple to detect the volatility structure predicted by the model in real-world longer-term interest rates. Second, the theoretical interest rate model is based on a number of assumptions which may not be satisfied for interest rates of all terms. For example, the model assumes that international investors consider Greek and Euroland-bonds (of the same term) as perfect substitutes and neglects risk-premiums. It should be emphasized here that the theoretical models of this paper are designed to capture the international links between the economies' financial markets on the road to a common currency, but do not account for microstructural or other institutional features on foreign exchange and capital markets. Theoretical extensions into these directions should be pursued in future research.

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Appendix

Table 1: Markov-switching AR(1)-GARCH(1,1) model for GRD-returns

	Regime 1 ($i = 1$) (low volatility)	Regime 2 ($i = 2$) (high volatility)
Parameter:		
a_{0i}	0.0008	0.0037
	(0.0051)	(0.0015)
a_{1i}	0.0324	−0.0070
	(0.0822)	(0.0939)
b_{0i}	0.0038	0.0011
	(0.0013)	(0.0002)
b_{1i}	0.4645	0.2352
	(0.1495)	(0.1131)
b_{2i}	0.6423	0.8201
	(0.0876)	(0.2370)
ν_i	2.4125	2.0668
	(0.4137)	(0.0565)
Transition probabilities:		
π_i	0.9977	0.9992
	(0.0037)	(0.0053)
Log-likelihood	1089.8923	
LB_1^2	0.0000	(0.9962)
LB_2^2	0.0126	(0.9937)
LB_3^2	0.0683	(0.9953)
LB_5^2	1.1139	(0.9529)
LB_{10}^2	1.2167	(0.9996)

Note: Estimates for parameters from the Eqs. (16) to (22). Standard errors are in parenthesis. LB_i^2 denotes the Ljung-Box-Q-statistic for serial correlation of the squared standardized residuals out to lag i. p-values are in parenthesis.

Table 2: AR(6)-GARCH(1, 1) model for 1-day (overnight) interest rate differentials

Parameter	Coefficient	Standard Error	z-Statistic	p-value
a_0	−0.0521	0.0280	−1.8594	0.0630
a_1	−0.1750	0.0737	−2.3743	0.0176
a_2	−0.3484	0.0729	−4.7799	0.0000
a_3	−0.2802	0.0742	−3.7768	0.0002
a_4	−0.2578	0.0836	−3.0852	0.0020
a_5	−0.1594	0.0867	−1.8396	0.0658
a_6	−0.1719	0.0682	−2.5190	0.0118
b_0	0.0092	0.0124	0.7411	0.4586
b_1 [ARCH(1)]	0.0860	0.0371	2.3181	0.0204
c_1 [GARCH(1)]	0.8612	0.1016	8.4759	0.0000
Log-Likelihood		100.0183		

Note: Estimates for parameters from the Eqs. (28) to (32).

Table 3: AR(20)-GARCH(1, 1) model for 1-month interest rate differentials

Parameter	Coefficient	Standard Error	z-Statistic	p-value
a_0	−0.0188	0.0072	−2.6129	0.0090
a_1	−0.0504	0.0629	−0.8026	0.4222
a_2	0.0344	0.0591	0.5819	0.5606
a_3	0.0774	0.0714	1.0845	0.2781
a_4	0.0350	0.0472	0.7425	0.4578
a_5	−0.0043	0.0512	−0.0834	0.9335
a_6	−0.0236	0.0573	−0.4127	0.6799
a_7	0.1395	0.0472	2.9549	0.0031
a_8	0.1192	0.0632	1.8865	0.0592
a_9	0.0731	0.0561	1.3021	0.1929
a_{10}	0.0889	0.0428	2.0771	0.0378
a_{11}	0.0243	0.0486	0.4994	0.6175
a_{12}	−0.1221	0.0529	−2.3077	0.0210
a_{13}	−0.1325	0.0415	−3.1918	0.0014
a_{14}	−0.0450	0.0381	−1.1819	0.2372
a_{15}	−0.0693	0.0432	−1.6063	0.1082
a_{16}	0.0132	0.0486	0.2722	0.7854
a_{17}	−0.0533	0.0618	−0.8622	0.3886
a_{18}	−0.1216	0.0485	−2.5054	0.0122
a_{19}	0.0401	0.0505	0.7947	0.4268
a_{20}	0.1710	0.0604	2.8313	0.0046
b_0	0.0001	0.0000	1.2029	0.2290
b_1 [ARCH(1)]	−0.0210	0.0355	−0.5910	0.5546
c_1 [GARCH(1)]	1.0073	0.0345	29.2152	0.0000
Log-Likelihood		248.9952		

Note: Estimates for parameters from the Eqs. (28) to (32).