

Gärtner, Dennis; Halbheer, Daniel

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Authors' addresses

Dennis Gärtner
E-mail: dennis.gaertner@soi.unizh.ch

Daniel Halbheer
E-mail: dhalbheer@soi.unizh.ch

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Sozialökonomisches Institut
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Rämistrasse 71
CH-8006 Zürich
Phone: +41-1-634 21 37
Fax: +41-1-634 49 82
URL: www.soi.unizh.ch
E-mail: soilib@soi.unizh.ch

Are There Waves in Merger Activity After All?*

Dennis Gärtner[†] and Daniel Halbheer

University of Zurich

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Abstract

This paper investigates the merger wave hypothesis for the US and the UK employing a Markov regime switching model. Using quarterly data covering the last thirty years, for the US, we identify the beginning of a merger wave in the mid 1990s but not the much-discussed 1980s merger wave. We argue that the latter finding can be ascribed to the refined methods of inference offered by the Gibbs sampling approach. As opposed to the US, mergers in the UK exhibit multiple waves, with activity surging in the early 1970s and the late 1980s.

Keywords: Merger Waves, Markov Regime Switching Regression Model, Gibbs Sampling

JEL Classification: G34, C32, C11, C15

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[†]Corresponding author. Address: University of Zurich, Socioeconomic Institute, Blümliisalpstr. 10, CH-8006 Zurich; tel.: +41 44 634 06 05; fax +41 44 634 49 07; e-mail: dennis.gaertner@soi.unizh.ch.

1 Introduction

There is broad consensus that mergers occur in waves. Since the seminal paper by Nelson (1959), many studies have reported a wave-like pattern in merger activity, pointing out the merger waves of the mid 1980s and mid 1990s in the US in particular.¹ Guided by these observations, a vast empirical literature has sought to identify potential causes and triggers for merger waves.² This empirical strand has more recently been complemented by efforts to explain the phenomenon of merger waves in the theoretical literature.³

While the general notion of mergers occurring in waves is practically undisputed, there is no clear consensus on how to operationalize the concept of a ‘merger wave’ in a time series context. The empirical literature has put forward three distinct approaches to modeling and identifying such waves. First, Golbe and White (1993) have sought to identify waves by fitting a sine curve to historic merger data. Second, merger series have been modeled by autoregressive processes capable of producing wave-like behavior (Shughart and Tollison, 1984; Clark et al., 1988; Chowdhury, 1993; Barkoulas et al., 2001). Third, and finally, merger series have been modeled by means of parameter-switching models where waves in activity are caused by discrete parameter switches (Town, 1992; Linn and Zhu, 1997).

This paper reexamines the case for detecting waves in merger activity in a time series context using more recent, consistent data and refined estimation techniques. Following Town (1992) and Linn and Zhu (1997), we employ a Markov regime switching model to describe the stochastic behavior of merger activity. We provide a thorough motivation for this approach, starting from Nelson’s (1959, p. 126) observation that aggregate merger series are characterized by “large bursts of activity separated by lengthy intervals of very low activity,” which we take to suggest the presence of two distinct unobserved states of merger activity, ‘high’ and ‘low’. By letting mean and variance of the autoregressive model be determined by realizations of the

¹Studies discussing the 1980s merger wave include Ravenscraft (1987), Golbe and White (1988, 1993), and Mitchell and Mulherin (1996). Mueller (1997), Andrade et al. (2001), and Harford (2005) describe the 1990s merger wave. An extensive review of earlier merger waves is provided in Scherer and Ross (1990, pp. 154–159).

²See e.g. Ravenscraft (1987), Shleifer and Vishny (1990), and Holmstrom and Kaplan (2001). Gugler et al. (2005) examine hypothesis that have been put forward as explanations of merger waves.

³Examples include Fauli-Oller (2000), Jovanovic and Rousseau (2002), Rhodes-Kropf and Viswanathan (2004), and Toxvaerd (2004).

Markov process governing the evolution of the two states, waves are triggered by switches in the unobserved state. While this approach borrows from Town (1992) and Linn and Zhu (1997), we propose a slightly modified formal specification in which the autoregressive processes' inertia persists also *across* state switches, leading merger activity to react less abruptly to such switches. More importantly, we use new and consistent quarterly time series data covering merger activity both in the US and the UK, extending from 1973:IV through 2003:IV and from 1969:I through 2003:IV, respectively.⁴

In this paper, we challenge the notion of the much-discussed 1980s merger wave in the US. We argue that the discrepancy between our findings and previous econometric identifications of this wave is driven by a further distinguishing feature of our analysis: the use of more recent estimation techniques. To address the central issue of wave identification, we conduct inference on the regime indicator within a Bayesian framework employing Gibbs sampling techniques (Gelfand and Smith, 1990; Albert and Chib, 1993). In contrast, the aforementioned studies by Town (1992) and Linn and Zhu (1997) base wave identification on Maximum Likelihood techniques (Hamilton, 1989, 1993). In this latter approach, inference consists in first estimating the model's unknown parameters via Maximum Likelihood, and then conducting inference on the unobserved state conditional on the parameter estimates. Bayesian analysis, on the other hand, avoids this two-step procedure by treating both the model parameters and state variable as random variables and basing inference on states on a joint distribution of parameters and states rather than on a conditional distribution. This methodological difference can lead to quite different conclusions regarding the likely path of the unobserved regime indicator if parameter uncertainty is sufficiently high, as the uncertainty on parameter estimates does not feed into uncertainty on states when employing a two-step estimation procedure.

Our main results are as follows. *First*, we find that the US have witnessed only the beginning of a wave in merger activity, this wave starting in 1995:IV. This result is consistent with the observations in Mueller (1997), Holmstrom and Kaplan (2001), and Andrade et al. (2001), all of which report an upsurge in merger activity in the mid 1990s. However, our investigation of industry level data does not support the

⁴Previous empirical studies examined the merger wave hypothesis using an assemblage of separate series differing in coverage and inclusion criteria. For a general discussion of available historical time series merger data and their limitations, see e.g. Golbe and White (1988).

prominent notion that waves in aggregate merger activity represent the clustering of surges within one or a few industries (Mitchell and Mulherin, 1996; Mulherin and Boone, 2000; Andrade et al., 2001; Harford, 2005). *Second*, even when fitting the model only to the data prior to the estimated break date, we fail to identify the much discussed 1980s merger wave. To explain our difference in findings, we argue that if there is sufficient uncertainty surrounding the model’s parameters, then the two-step Maximum Likelihood estimation procedure can convey a deceptive degree of certainty about state inference. *Third*, the UK has witnessed two merger waves, the first starting in 1971:I and ending in 1973:IV and the second lasting from 1986:III to 1989:IV. The dating of these merger waves is close to the evidence reported in Hughes (1993, p. 16).

The remainder of the paper is organized as follows. Section 2 briefly comments on the data employed. Section 3 provides a thorough motivation of the model. In Section 4, we describe the inference problem and give a brief introduction to the Gibbs sampling approach. Section 5 presents the main results of our estimation both for the US and the UK series. Section 6 concludes.

2 The Data

Our paper follows the majority of previous studies, particularly Town (1992) and Linn and Zhu (1997), in using the number of transactions as the measure of historical merger activity.⁵ Specifically, we investigate the following series:

- (i) *The US merger series, covering 1973:I–2003:IV*. The time series data are taken from various issues of *Mergerstat Review*, a publication by FactSet Mergerstat LLC. The series reports publicly announced mergers, acquisitions and unit divestitures involving (i) at least one US company, (ii) a transaction volume exceeding \$1 million, and (iii) a purchase price exceeding 10% of the acquired company’s equity (i.e., an interest exceeding 10% of the acquired firm’s equity).
- (ii) *The UK merger series, covering 1969:I–2003:IV*. These data are published by the Office for National Statistics on a quarterly basis. The series consists of

⁵Other prominent measures of merger activity suggested in the literature are the dollar value of merger transactions (see e.g. Golbe and White, 1988, Scherer and Ross (1990) and the number of transactions relative to appropriate population totals (see e.g. Hughes, 1993, p. 16; Gugler et al., 2005).

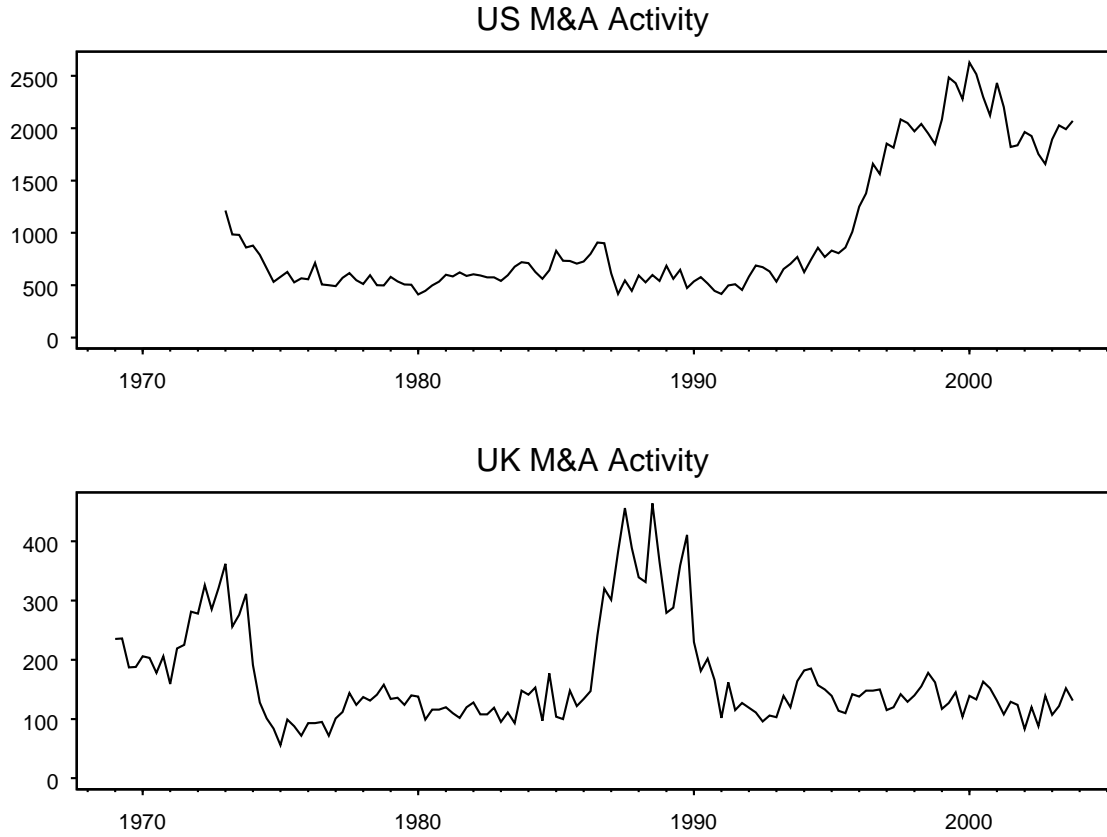


Figure 1: The Number of Merger Transactions, US Series, 1973:I–2003:IV (in Top Panel) and UK Series, 1969:I–2003:IV (Bottom Panel).

publicly announced mergers and acquisitions involving UK companies only. In contrast to the US data, there is no explicit cut-off bias relating to the value of the transaction, but the deal has to aim at gaining de jure control of the acquired company (i.e., a controlling interest exceeding 50% of the acquired firm’s equity).

Plots of these series are presented in Figure 1.

3 A Markov-Switching Model of Merger Waves

As outlined above, the literature has advanced the idea that mergers follow a wave pattern. We take this casual impression to suggest the presence of two distinct states

of merger activity, high and low, as follows:⁶

Assumption 1. *Each period t is associated with an unobserved latent state variable $S_t \in \{1, 2\}$, where $S_t = 1$ implies that period t is a low-activity period and $S_t = 2$ denotes a high-activity period.*

The basic idea is then to let unobserved switches between states of high and low activity feed into observed merger activity—in a sense to be made precise shortly—so as to induce the alleged wave-like behavior. Hence, given Assumption 1, the remaining key questions concerning our description of mergers are: (1) What determines the unobservable state S_t in any period t , and (2) how exactly do the unobservable states feed into observed merger activity y_t ?

The general framework in which we deal with these questions is the Markov regime switching model originally proposed by Hamilton (1989). In a nutshell, this approach treats both the sequence of observations y_t and the sequence of states S_t as (interdependent) random variables, specifies a model which jointly generates the two sequences, and then estimates the model using the observed series y_t while treating the sequence of states as ‘missing data’. This framework offers several advantages over more traditional approaches to break-point analysis which typically rely on casual determination of candidate break-dates or *ad hoc* restrictions on the number of break dates (see e.g. Chow, 1960, 1984; Andrews, 1993): First and foremost, a major goal of our analysis is not only the estimation of regime-dependent structural model parameters, but dating the waves (i.e. conducting inference on the break dates themselves). This in turn requires modelling the probability law governing changes in regime rather than imposing particular break-dates *a priori*. Through the probability model, we can then let the data itself speak about the likely incidence of significant changes. Second, we would like to propose a unified structural process capable of describing various merger series (such as across countries or industries) with apparently different frequency and timing of waves, which requires that wave dates be determined endogenously by the process.

More specifically, concerning the determination of states, we shall assume that

⁶We shall comment on the idea of using more than two states when discussing our estimation results further below. Let us just note for now that raising the number of attainable states invokes the usual trade-off between achieving a better fit to the data and overparameterizing the model. As a consequence, we suggest using the minimal number of states capable of producing the described behavior in mergers.

states follow an independent first-order Markov process. Thus, in any period t , the probability of switching to a certain state in the next period $t + 1$ depends only on the state in period t . Specifically, we assume the following:

Assumption 2. *The unobserved state variable S_t follows a first-order Markov process with transition probabilities from any period t to period $t + 1$ given by*

$$\Pr(S_{t+1} = 1|S_t = 1) = p_{11} \quad \text{and} \quad \Pr(S_{t+1} = 2|S_t = 2) = p_{22}, \quad (1)$$

with $p_{11}, p_{22} \in [0, 1]$. In any period t , these transition probabilities are independent of past (log) merger realizations (y_t, y_{t-1}, \dots) .

It is important to note that ‘merger waves’ as we understand and model them need not display a highly regular periodic pattern. Indeed, the first-order Markov specification implies that the process governing the states displays very little memory. This low-memory approach seems justified by the aforementioned literature giving little impression that the documented bursts of high activity display a highly regular periodic pattern.⁷ Some structure is of course nonetheless implied by our Markov specification, such as the expected duration of a high state being $p_{22}/(1-p_{22})$ and the expected duration of a low state being $p_{11}/(1-p_{11})$, but these durations generally display a rather high variability. Furthermore, due to the first order Markov property, the *remaining* expected duration of a certain state is independent of how long the process has already been in that state, which again reflects the low-memory quality of the process.

Finally, note that the Markovian model encompasses the extreme possibility of a state being ‘absorbing’ in the sense that, once the process reaches a certain state, it remains in that state indefinitely (so that the regime switch is permanent rather than transitory). This is the case for the low-activity state if $p_{11} = 1$ and for the high-activity state if $p_{22} = 1$. Conversely, whenever this is not the case, so $p_{11}, p_{22} < 1$ and if in addition $p_{11} + p_{22} > 0$ (so there is no completely deterministic alternation between states), then the Markov chain turns out to be *ergodic* (see e.g. Hamilton, 1989). Then, a further key characteristic of the state switching model is given by the ergodic regime probabilities $\Pr(S_t = i)$, i.e. the unconditional probability of state

⁷Although Golbe and White (1993) do report evidence of a sine wave pattern in US merger activity based on data up to 1989, by inspection of the plots in Figure 1 we strongly suspect that their model would no longer provide a very good fit to our more recent series.

$i \in \{1, 2\}$.⁸ These can be shown to be given by $\Pr(S_t = 1) = (1 - p_{11})/(2 - p_{11} - p_{22})$ and $\Pr(S_t = 2) = 1 - \Pr(S_t = 1)$.

Let us now turn to the second question concerning the exact form of the state's impact on merger activity. We start from the idea that merger activity follows some sort of mean reverting autoregressive process and augment this by assuming that both the mean and the variance of this process are time-varying and dependent on the states.⁹ Specifically, we make the following assumption:

Assumption 3. *Conditional on the sequence of unobserved states S_t , (log) mergers y_t follow the AR(k) process*

$$y_t - \mu_{S_t} = \sum_{i=1}^k \phi_i (y_{t-i} - \mu_{S_t}) + \varepsilon_t, \quad (2)$$

where (i) the ε_t are independently $N(0, \sigma_{S_t}^2)$ and independent of previous merger realizations $(y_{t-1}, y_{t-2}, \dots)$, (ii) $\mu_{S_t} \in \{\mu_1, \mu_2\}$ and $\sigma_{S_t} \in \{\sigma_1, \sigma_2\}$ are determined by the state in period t , (iii) $\mu_2 \geq \mu_1$, and (iv) the autoregressive coefficients $\phi_1, \phi_2, \dots, \phi_k$ are restricted so that the roots of the associated lag polynomial, $\phi(L) \equiv 1 - \phi_1 L - \phi_2 L^2 - \dots - \phi_k L^k$, lie outside the complex unit circle.

We let the idea that $S_t = 2$ entails higher activity impose the normalization $\mu_2 \geq \mu_1$, while σ_1 and σ_2 are left unrestricted (except for the obvious nonnegativity requirement). Furthermore, the familiar condition on the autoregressive parameters $\phi_1, \phi_2, \dots, \phi_k$ ensures that the process is in some sense mean reverting, where this mean however depends on the state sequence. Put somewhat differently, the condition ensures that the only source of non-stationarity is through switches in the regime (i.e. given a constant sequence of states, $S_t = i$ for all t and $i \in \{1, 2\}$, y_t follows a stationary process).¹⁰

⁸A convenient way to think of the ergodic probabilities is in terms of the fraction of high and low states observed in an infinitely long realization of the Markov chain.

⁹A previous study by Shughart and Tollison (1984) reports little success in describing waves in merger activity as a standard autoregressive process with *constant* mean and variance, $y_t - \mu = \sum_{i=1}^k \phi_i (y_{t-i} - \mu) + \varepsilon_t$ with $\varepsilon_t \sim \text{i.i.d. } N(0, \sigma^2)$, where the wave property would be reflected solely by some of the higher-order autoregressive coefficients ϕ_2, \dots, ϕ_k being nonzero. However, such a specification can produce only rather ‘tame’, linear wave-like oscillations, while we suspect that the large bursts of activity separated by long intervals of low activity identified in the aforementioned literature can only be reconciled with a nonlinear model such as ours.

¹⁰The literature has also proposed non-mean reverting processes such as random walks to describe merger activity (see e.g. Chowdhury, 1993). Even though standard tests reject the unit root

Specification (2) differs in a small but important way from Hamilton's (1989) original specification,

$$y_t - \mu_{S_t} = \sum_{i=1}^k \phi_i (y_{t-i} - \mu_{S_{t-i}}) + \varepsilon_t, \quad (3)$$

which has been the workhorse model in the literature on mean and variance switching Markov models and also happens to be the model used by Town (1992) and Linn and Zhu (1997) to describe mergers in particular. The subtle but important difference is that specification (2) assumes 'sluggish' adjustments of the merger series to a state switch, whereas by specification (3), state switches cause an immediate full shift in activity. To see this, observe that in specification (2), what systematically affects today's deviation from the mean, $y_t - \mu_{S_t}$, is a weighted sum of past deviations from the *current* mean, $y_{t-i} - \mu_{S_t}$, whereas in specification (3), today's deviation is determined by past deviations from the *contemporaneous* mean, $y_{t-i} - \mu_{S_{t-i}}$. Thus, the two models imply rather different dynamic consequences of a shift in regime. This is most effectively illustrated by setting $k = 1$ in both (2) and (3) above and considering a permanent shift from state 1 to state 2 between dates t and $t + 1$. According to specification (3), the switch to state 2 at date t raises the value of any subsequent y_{t+j} ($j \geq 0$) by $\mu_2 - \mu_1$ over its respective value if no state-switch had occurred. In model (2), on the other hand, the impact of the state switch at t only raises subsequent y_{t+j} by $(1 - \phi_1^j)(\mu_2 - \mu_1) \leq \mu_2 - \mu_1$ for any $j \geq 0$.¹¹ Hence, model (3) suggests that the merger series immediately jumps toward the new mean after a state switch, whereas model (2) describes a more gradual, 'sluggish' gravitation toward the new mean. Note however that the difference of the state switch's impact between the models disappears as j rises, so that the models differ most markedly during the adjustment period.

We favor specification (2) over (3) for two somewhat interrelated reasons. First, casual inspection of real merger data suggests that the transition to a significantly higher (or lower) level of merger activity is indeed sluggish rather than immediate.

hypothesis for our UK merger series, this is indeed not the case for the US series. However, it is a well understood fact that in general, unit-root tests have very little power over Markov-switching alternatives (see e.g. Nelson et al., 2001), so that such tests do not invalidate our proposed model. Furthermore, if we perform the unit root tests using only US data prior to 1995:III (which amounts to discarding little more than a quarter of the data), the unit root hypothesis is clearly rejected. We take this as evidence favoring our Markov-switching model over the random walk hypothesis.

¹¹Recall that in an AR(1) model, $|\phi_1| < 1$ by the restriction on the autoregressive coefficients in Assumption 3.

Second, perhaps contrary to other common applications of mean switching models, there seems to be no intuitive reason to suggest that the merger process does not display the same amount of inertia when switching to a high or low activity state as within a given state. Indeed, if for instance we suspect the sluggishness in merger series to be a consequence of the fact that real world mergers may take considerable time to process (due to preparation, approval, etc.), thereby causing sluggish adjustment to any unobserved structural shocks, then this sluggishness should persist also when the economy moves to a generally higher or lower level of activity (i.e. when it is hit by a ‘large’ shock).¹² For these reasons, we shall employ specification (2) for the remainder of our analysis.¹³

As a final remark, we should point out that more generally, mean and variance switching is not the only way in which high and low activity states may be thought to affect mergers. For instance, an alternative specification might have states impact only the growth rate rather than the mean level of the merger series.¹⁴ However, we view the mean-switching specification as closest in spirit to the wave notion developed in the literature. What may nonetheless seem somewhat extreme at first sight is that our mean switching model appears to posit that waves always have the same magnitude (or, stated differently, the described ‘bursts of activity’ always have the same magnitude).¹⁵ However, we would like to argue that empirically, a major task in identifying waves is being able to tell actual waves from smaller ‘ripples’, and a straightforward way to accomplish this is to posit that waves always have a certain height. We will return to this point in our discussion of the 80s merger wave in Section 5.3.

¹²This argument can be formalized by noting that specification (2) can be interpreted as a standard $AR(k)$ model where the state only affects the distribution of the error term. This can be seen by rewriting model (2) as $y_t = \sum_{i=1}^k \phi_i y_{t-i} + \tilde{\varepsilon}_t$, $\tilde{\varepsilon}_t \sim N(\tilde{\mu}_{S_t}, \sigma_{S_t}^2)$, where $\tilde{\mu}_{S_t} = (1 - \sum_{i=1}^k \phi_i) \mu_{S_t}$.

¹³A nice technical side-effect of using specification (2) is that inference on states does not involve an approximation (see e.g. Kim and Nelson, 1999, pp. 68–70).

¹⁴For instance, Owen (2004) pursues such an idea for UK mergers by fitting a three-state mean switching Markov model to the differenced level data $\Delta y_t \equiv y_t - y_{t-1}$.

¹⁵Note that in our model with sluggish adjustment to the mean, this is only strictly true in expectation for waves having the same duration. If the process is sluggish, then the shorter a wave, the lower its peak.

4 Estimation Techniques

We estimate the model parameters and the path of the latent Markov switching regime indicator within a Bayesian framework employing Markov chain Monte Carlo simulation methods. Letting $\boldsymbol{\beta} \equiv (\mu_1, \mu_2, \sigma_1, \sigma_2, \phi_1, \phi_2, \dots, \phi_k, p, q)$ denote the model's parameters, letting $\mathcal{Y}_T = (y_1, y_2, \dots, y_T)$ denote the data observed, and letting $\mathcal{S}_T \equiv (S_1, S_2, \dots, S_T)$ denote the unobserved sequence of states, Bayesian inference in our model takes the form of using the data \mathcal{Y}_T and the model specified in Section 3 to map a given prior distribution of parameters, $p(\boldsymbol{\beta})$, into a joint posterior distribution of states and parameters, $p(\mathcal{S}_T, \boldsymbol{\beta} | \mathcal{Y}_T)$.

Rather than investigating $p(\mathcal{S}_T, \boldsymbol{\beta} | \mathcal{Y}_T)$ analytically, Markov chain Monte Carlo methods provide a simple way of simulating draws from this distribution. We use a particular form of these methods, the Gibbs sampling technique, which is an iterative scheme based on simulating successive draws from the conditional posterior distributions of the state vector \mathcal{S}_T and the appropriately partitioned parameter vector $\boldsymbol{\beta}$:

- (i) $p(\mathcal{S}_T | \boldsymbol{\beta}, \mathcal{Y}_T)$
- (ii) $p(\mathbf{p} | \boldsymbol{\mu}, \boldsymbol{\sigma}, \boldsymbol{\phi}, \mathcal{S}_T, \mathcal{Y}_T)$
- (iii) $p(\boldsymbol{\mu} | \boldsymbol{\sigma}, \boldsymbol{\phi}, \mathbf{p}, \mathcal{S}_T, \mathcal{Y}_T)$
- (iv) $p(\boldsymbol{\sigma} | \boldsymbol{\mu}, \boldsymbol{\phi}, \mathbf{p}, \mathcal{S}_T, \mathcal{Y}_T)$
- (v) $p(\boldsymbol{\phi} | \boldsymbol{\mu}, \boldsymbol{\sigma}, \mathbf{p}, \mathcal{S}_T, \mathcal{Y}_T)$,

where $\boldsymbol{\mu} \equiv (\mu_1, \mu_2)$, $\boldsymbol{\sigma} \equiv (\sigma_1, \sigma_2)$, $\boldsymbol{\phi} \equiv (\phi_1, \phi_2, \dots, \phi_k)$, and $\mathbf{p} \equiv (p_{11}, p_{22})$.¹⁶

In contrast to the *full* posterior $p(\mathcal{S}_T, \boldsymbol{\beta} | \mathcal{Y}_T)$, each of the *marginal* posterior distributions (i)–(v) can be handled analytically. Simulated draws from (i)–(v) are thus easily obtained, and the Gibbs sampler provides a way of iterating on such draws to simulate draws from the full posterior $p(\mathcal{S}_T, \boldsymbol{\beta} | \mathcal{Y}_T)$.¹⁷

Given the tool to generate representative sets of draws from $p(\mathcal{S}_T, \boldsymbol{\beta} | \mathcal{Y}_T)$, properties of this distribution such as individual parameters' marginal distributions

¹⁶For a general introduction to Gibbs sampling, readers are referred to Geman and Geman (1984) and Gelfand and Smith (1990). A textbook treatment of the method can also be found in Kim and Nelson (1999).

¹⁷The derivation of these posterior distributions is standard and available from the authors upon request.

and moments are easily characterized by use of their sample equivalents. To address the central issue of wave identification, we will be particularly interested in $\Pr(S_t = 2|\mathcal{Y}_T)$, the posterior probability of being in a high merger activity state at any date t . These probabilities are obtained by averaging across the simulated paths for the states, each simulated while cycling through the above posterior distributions.

Basing state inference on $\Pr(S_t = 2|\mathcal{Y}_T)$, while natural in the Bayesian framework of Gibbs Sampling, elegantly overcomes a potential pitfall to classical inference. In a classical setting, inference on states is obtained through a two-step procedure by first obtaining a Maximum Likelihood parameter estimate $\hat{\beta}$, and then calculating $\Pr(S_t = 2|\mathcal{Y}_T, \hat{\beta})$, the probability of $S_t = 2$ in any period t under the assumption that $\hat{\beta}$ corresponds to the *true* parameter values.¹⁸ From a Bayesian perspective, the derived inference on states is thus to be read as contingent on the econometrician having full confidence in his parameter estimate $\hat{\beta}$. But only rarely will this correspond to the econometrician's true confidence in $\hat{\beta}$. Moreover, alternative conceivable values of β will typically lead to different values of $\Pr(S_t = 2|\mathcal{Y}_T, \beta)$, so that uncertainty about β will feed into uncertainty on states. As a result, basing state inference on $\Pr(S_t = 2|\mathcal{Y}_T, \hat{\beta})$ rather than on $\Pr(S_t = 2|\mathcal{Y}_T)$ can convey a false degree of certainty about states by neglecting uncertainty about parameters.

To make this important point more transparent, note that $\Pr(S_t = 2|\mathcal{Y}_T)$ and $\Pr(S_t = 2|\mathcal{Y}_T, \hat{\beta})$ are related by

$$\Pr(S_t = 2|\mathcal{Y}_T) = \int \Pr(S_t = 2|\mathcal{Y}_T; \beta) p(\beta|\mathcal{Y}_T) d\beta, \quad (4)$$

where $p(\beta|\mathcal{Y}_T)$ denotes the posterior density of the parameter vector. By (4), our Bayesian inference based on $\Pr(S_t = 2|\mathcal{Y}_T)$ can be interpreted as considering $\Pr(S_t = 2|\mathcal{Y}_T; \beta)$ for *any* conceivable parameter constellation β —that is, for the maximum likelihood estimate $\hat{\beta}$ in particular, but also for any other conceivable β —, weighting it with the respective posterior probability of β , and adding up across β to produce $\Pr(S_t = 2|\mathcal{Y}_T)$. It is straightforward to see from (4) that, if posterior parameter uncertainty is sufficiently high and conditional inference on S_t is sufficiently sensitive to β , then $\Pr(S_t = 2|\mathcal{Y}_T, \hat{\beta})$ can differ substantially from $\Pr(S_t = 2|\mathcal{Y}_T)$. We will provide an impressive illustration of this difference below in our investigation of US

¹⁸For details on standard maximum likelihood methods, refer to Hamilton (1989, 1993).

merger activity, where we shall in fact argue that this methodological point is likely to have played a substantial role in previous identifications of a 1980s merger wave.

5 Estimation Results

This section reports the results of fitting our lagged mean and variance switching model as given by Assumptions 1 through 3 to US and UK log merger series.¹⁹ In a first step, Section 5.1 analyzes the entire quarterly US merger series from 1973 through 2003. Its main finding, the identification of a wave beginning in the mid 1990s, is augmented by a look at industry level data in Section 5.2. In a second step, to investigate the lack of evidence for an 1980s merger wave in more detail, Section 5.3 reestimates the model using only data up to 1995. Section 5.4 argues that in so doing, the estimation techniques described in Section 4 play a decisive role. Finally, Section 5.5 offers a shift of focus to the UK by fitting the model to the quarterly UK merger series spanning 1969 through 2003.

As outlined, inference is conducted in a Bayesian context using the Gibbs sampling method to derive posterior distributions of the model parameters and to assess the sequence of unobserved states. In all cases, Gibbs sampling involved 20,000 iterations, where a burn-in sequence consisting of the first 1,000 draws was discarded prior to any inference. Output of the sampler was closely monitored to ensure proper convergence of the filter.

5.1 The Tidal Wave of the Mid 1990s:

US Mergers, Full Series 1973:I through 2003:IV

First, we consider the full series on US mergers over the entire available time span from 1973:I through 2003:IV, as presented in Figure 1. Preliminary estimation of the model with various lag lengths k suggested setting $k = 4$.

¹⁹The main reason for using log rather than level merger data is that all series considered are nonnegative by construction. Strictly speaking, the model as defined by (1) and (2) therefore provides no valid description of the level series due to its capability of producing negative observations. However, all subsequent inference was nonetheless also conducted after fitting the level merger data to the model. Qualitative results, specifically concerning inference on states, differ only little from those obtained for the log merger series. Any remaining noteworthy differences are explicitly pointed out in the subsequent discussion.

	priors		posteriors				ML
	mean	std	mean	median	std	95%-band	MLE
μ_1	6.7	1000.00	6.386	6.390	0.083	(6.204, 6.538)	6.3959
μ_2	6.7	1000.00	7.687	7.677	0.103	(7.510, 7.924)	7.6653
σ_1	0.2	0.28	0.137	0.136	0.008	(0.123, 0.153)	0.1320
σ_2	0.2	0.28	0.096	0.095	0.010	(0.079, 0.118)	0.0834
ϕ_1	0.0	1.00	0.683	0.683	0.098	(0.488, 0.874)	0.6578
ϕ_2	0.0	1.00	0.063	0.063	0.113	(−0.159, 0.287)	0.0538
ϕ_3	0.0	1.00	−0.114	−0.114	0.112	(−0.334, 0.103)	−0.1136
ϕ_4	0.0	1.00	0.146	0.145	0.085	(−0.021, 0.313)	0.1386
p_{11}	0.9	0.21	0.987	0.991	0.012	(0.955, 1.000)	0.9886
p_{22}	0.9	0.21	0.997	1.000	0.010	(0.969, 1.000)	1.0000

Note: Result of 20,000 Gibbs-Sampling iterations, iterations 1,000 through 20,000 used for inference. 95%-band refers to 95% posterior probability bands.

Table 1: Estimation Results for US Merger Activity, 1973:I through 2003:IV.

For all model parameters, Table 1 gives summary statistics both on the priors used and on the marginal posterior distributions obtained. The priors on all parameters were chosen to be relatively uninformative within the class of admissible conjugate priors (which are: normal distributions for the parameters μ_{S_t} and ϕ_i , inverted gamma distributions for $\sigma_{S_t}^2$, and beta distributions for p_{11} and p_{22}). To give an impression of the posterior parameter distributions beyond the mere summary statistics, Figure 2 plots histograms representing the estimated marginal posterior distributions of the parameters. Despite our focus on Bayesian inference, Table 1 also gives maximum likelihood point estimates of all parameters in the last column.

The first feature obvious from Table 1 is that the data indeed leads to significant updates in the priors on all parameters, as shown by direct comparison of the standard deviations of priors and posteriors. Furthermore, Figure 2 reveals that the corresponding marginal posterior distributions are single peaked and well behaved. As may have been expected from a glance at the original data, comparing the posteriors for μ_1 and μ_2 reveals that mean log merger activity in the high state 2 significantly exceeds that in the low state 1.

Although not literally interpretable as state-contingent means of the untransformed series due to the nonlinear log transformation, the median values of $\exp(\mu_1)$

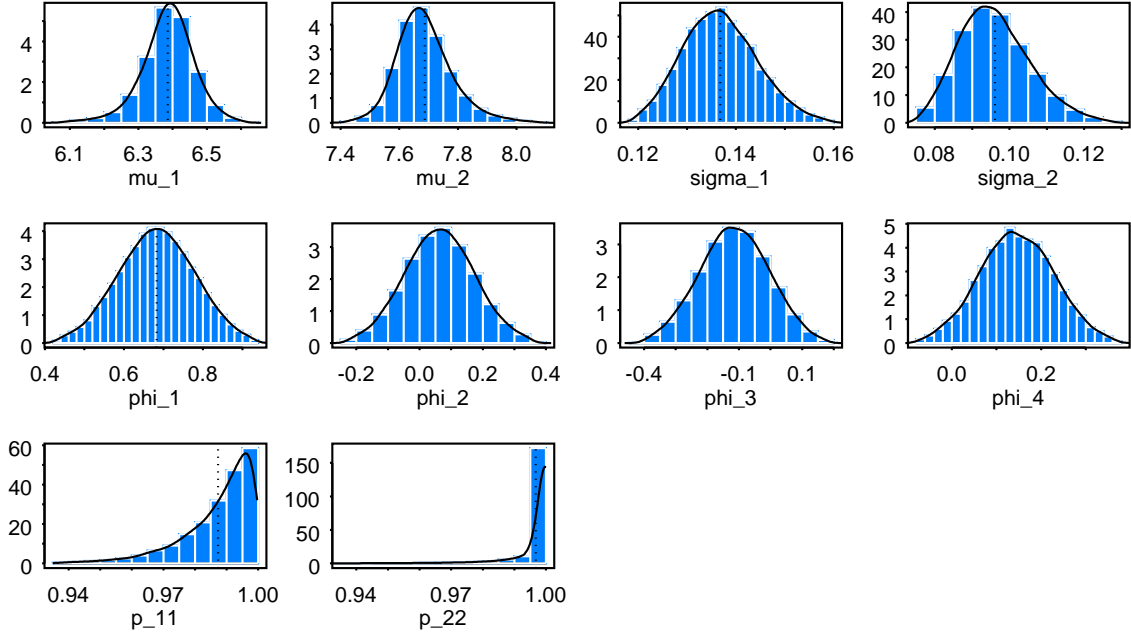


Figure 2: Marginal Posterior Distributions of Parameters for US Log Merger Series 1973:I–2003:IV.

and $\exp(\mu_2)$ suggest that in level terms, merger activity in the low and high activity state are in the region of 596 and 2,160 mergers per quarter, respectively. Furthermore, log mergers seem to be significantly less volatile in the high activity state, as shown by a comparison of the posteriors on σ_1 and σ_2 . Again however, caution is called for when drawing conclusions concerning the volatility of the *level* merger series, as the log transformation compresses differences at higher absolute levels of activity. Indeed, fitting the model to the level merger data suggests that in level terms, mergers are significantly more volatile in the high-activity state. Next, estimates on the autoregressive coefficients ϕ_i show that mergers display a considerable degree of inertia also *within* states. Moreover, mean and median of the largest root across samples both figure at 0.84, suggesting a significant degree of autocorrelation in the merger series. Finally, the posteriors on the transition probabilities p_{11} and p_{22} let us conclude that switches in regime are rather unlikely. Specifically, the median expected duration of a low activity state, $p_{11}/(1 - p_{11})$ is approximately 27 years (mean expected duration is heavily influenced by the skewness in the posterior on q and lies around 133 years). Conversely, estimates on p_{22} show that the high activity state seems to be essentially absorbing, so that a regime of high merger activity is

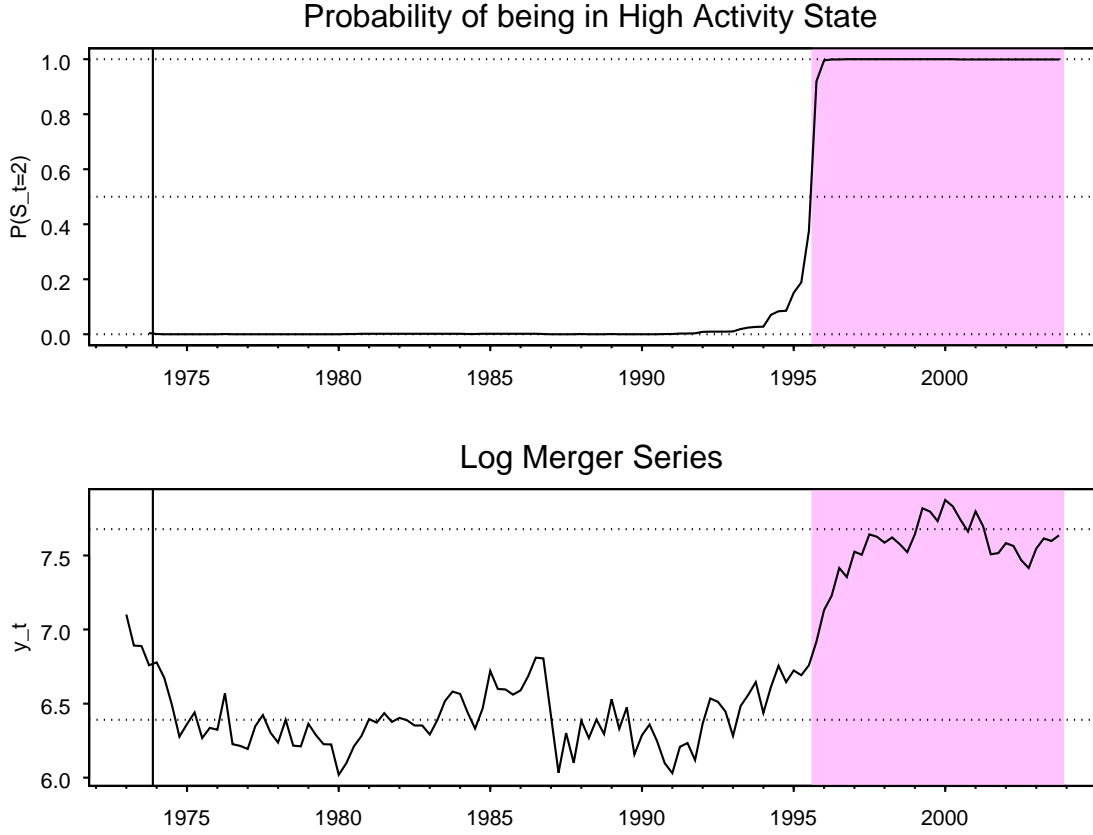


Figure 3: Estimated Probability of Being in State of High Merger Activity, US Log Merger Series 1973:I–2003:IV (in Top Panel, Bottom Panel displays Log Merger Series).

unlikely ever to be left again—this result being driven, of course, by the fact that there does not seem to have been a single reversion from high to low activity in the series so far. Finally, for the ergodic probability of being in a high state (unconditional on the data), $\Pr(S_t = 2) = (1 - p_{22}) / (2 - p_{11} - p_{22})$, the mean posterior is 86.5%, whereas the median is 99.7%.

With these results on the model’s parameters in mind, let us now return to our main objective, the identification of waves in mergers. Figure 3 plots the probability of being in a high state in any period t given the observed US merger data, $\Pr(S_t = 2 | \mathcal{Y}_T)$. This probability plot is shown in the top panel of Figure 3, whereas the bottom panel reproduces the underlying log merger series (along with the posterior median of μ_1 and μ_2 as dashed horizontal lines) for convenience. To highlight the most likely state for any period t , periods for which $\Pr(S_t = 2 | \mathcal{Y}_T) > 0.5$ are shaded. However, any interpretation of this ‘best guess’ should take account of the underlying

value of $\Pr(S_t = 2|\mathcal{Y}_T)$ as a straightforward measure of confidence in this guess: The closer $\Pr(S_t = 2|\mathcal{Y}_T)$ to 0.5, the more uncertainty surrounds the best guess.

Figure 3 shows that our estimation produces strikingly clear inference concerning the unobserved state. First, we find compelling evidence that over the entire period between 1973 and 2003, US merger activity has in fact experienced only a single regime switch, that switch being from low to high activity. Cast into the wave terminology, this suggests that since 1973, the US have so far witnessed only the beginning of a single ‘tidal wave’ in merger activity. Second, while this observation alone may come as no major surprise given a glance at the log merger plot, the clear-cut jump in the probability plot also allows us to date the beginning of this tidal wave rather precisely. Specifically, the assessed probability of a high state of merger activity jumps from 0.189 in 1995:II and 0.375 in 1995:III to a value of 0.920 in 1995:IV. We may thus conclude that the wave in US merger activity is very likely to have been triggered between the third and fourth quarter of 1995.²⁰

Rounding up, we should stress three points relevant to the interpretation of these results. First, as pointed out in Section 4, using $\Pr(S_t = 2|\mathcal{Y}_T)$ —rather than $\Pr(S_t = 2|\mathcal{Y}_T, \hat{\beta})$ for a point estimate of the parameters $\hat{\beta}$ —means taking account of uncertainty about the model’s structural parameters for the inference on states. It is all the more noteworthy that Figure 3 conveys an appreciably clear message concerning the likely sequence of states. Second, although we shall more thoroughly investigate the failure to identify the hump in merger activity around the mid 1980s as a wave in a moment in Section 5.3, let us note for now that this result is even more clear-cut when fitting the state-switching model to the level rather than the log-merger data. Intuitively, this is due to the simple fact that the log transformation exaggerates the mid 1980s hump in merger activity relative to the level data. Finally, we should point out our estimation procedure’s weakness in producing inference on states at the very beginning of the series: For technical reasons, inference on states actually begins only in 1974:I (rather than in 1973:I), and inference on states in these early periods is generally somewhat sensitive to starting values used.²¹

²⁰Specifically, the probability of the US having witnessed a *single* state switch from low to high between 1995:III and 1995:IV (rather than at any other date) can be calculated at 54.2%, which is contrasted by the probability of a corresponding single switch one quarter earlier (i.e. between 1995:II and 1995:III) of 18.5%, of 3.7% two quarters earlier, and 7.5% one quarter later.

²¹To understand the first point, note that our estimation of an $\text{AR}(k)$ model takes the first k observations of y_t as exogenously given, which is why inference on states begins in 1974:I. Concerning the second point, inferring states by means of the Hamilton filter requires specifying

5.2 Reflections of the US 1990s Wave at the Industry Level

As a prominent explanation of the clustering of merger activity in time, the literature has advanced the idea that surges in aggregate merger activity represent firms' optimal responses to industry-level shocks.²² According to this hypothesis, waves in merger activity at the aggregate level will be the result of temporary surges in merger activity in one or a few industries. Concerning our above findings, this naturally raises the question of whether the marked increase in US merger activity in the mid 1990s was dominated by any specific industries. While the available data do not permit us to estimate our model at the industry level, casual investigation of annual industry level data suggests that the mid 1990s wave is hardly attributable to one or a few industries alone.

To this end, of the 50 industries identified by the *Mergerstat Review*, Figure 4 plots *annual* merger data for those eleven US industries with the strongest merger activity between 1990 and 2003, where industries were ranked according to overall activity (in terms of numbers of mergers) for that period. While Figure 4 shows that industries were certainly not uniformly hit by a wave in 1995, it is nevertheless apparent that the resulting aggregate wave is anything but the result of a single industry level burst. For instance, the largest industry level share in overall annual merger activity was in Computer Software, Supplies and Services at its pronounced peak in 2000, with a share of 26%. Although significant, such shares still make it impossible to explain the pronounced jump in aggregate mergers—from around 600 per quarter before the wave to over 2,000 thereafter—as caused only by a small subset of industries. Furthermore, as the bottom right plot in Figure 4 shows, even after removing the 11 most active industries (which account for 52% to 66% of annual merger activity between 1990 and 2003), residual merger activity still gives

the initial exogenous probability of being in a high activity state, $\Pr(S_0 = 2|\beta; y_{-k+1}, \dots, y_0)$. All inference shown was produced using an uninformative but nonetheless somewhat arbitrary initial probability of 0.5, although simulations with alternative initial probabilities were run to check the results for robustness. Specifically, for the full US merger series considered here, results turn out to be very robust despite the relatively high level of merger activity at the very beginning of the series. Intuitively, this stems from the fact that, even though the data may suggest that the US may have been in a high level of merger activity immediately prior to 1973, the strong downward trend at the beginning of the sampling period clearly leads us to conclude that the US economy must nevertheless have already found itself in a low state of merger activity at the time our inference on states begins.

²²See for instance Mitchell and Mulherin (1996), Mulherin and Boone (2000), Andrade et al. (2001), and Harford (2005).

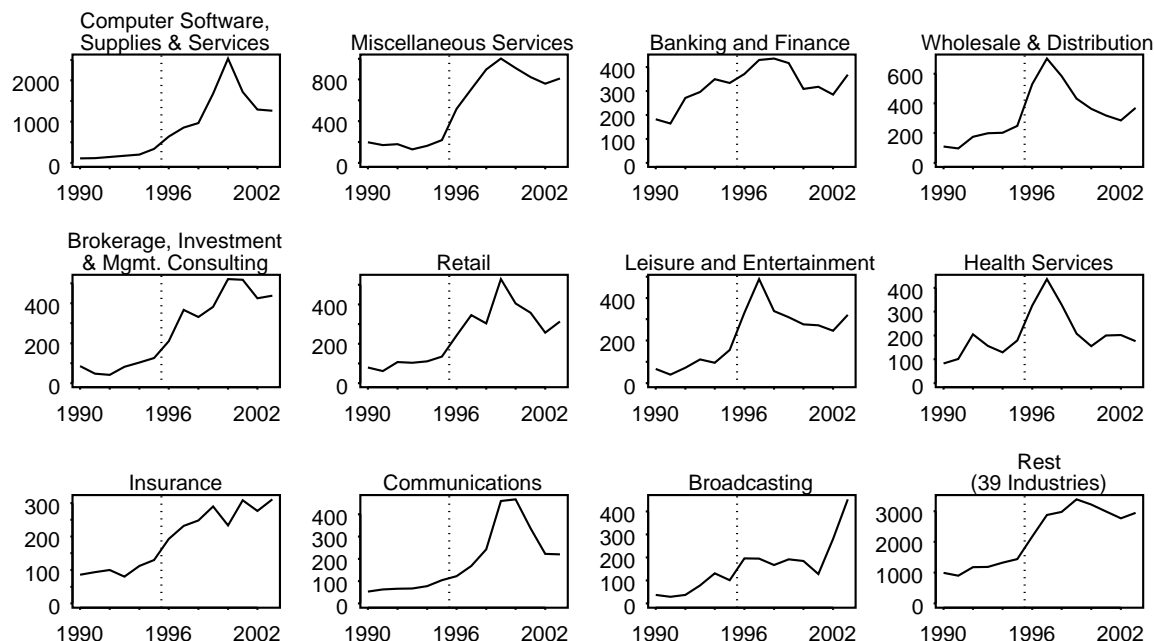


Figure 4: Annual US Industry Level Merger Activity, 1990 through 2003 (Source: *Mergerstat Review*).

a strong impression of a mid 1990s merger wave.

5.3 What ever happened to the 1980s Merger Wave?

US Mergers 1973:I through 1995:III

One of the most striking findings in Section 5.1 has been our failure to identify even the faintest hint of the much-discussed 1980s merger wave. Indeed, a simple look at the log merger series depicted in Figure 3 may indeed raise questions about there being something less pronounced—but nonetheless ‘wave-like’—about the visible hump in merger activity in the mid 1980s.

As pointed out previously, an arguably rather strong assumption implicit in our two-state model is that waves always come in similar sizes (where ‘size’ refers to a wave’s peak height). Along these lines, a perfectly valid reservation with our result might be that, even if US merger activity between 1973 and 2003 was dominated by a single gigantic tidal wave in the mid 1990s, the assumption of similarly-sized waves downplays the importance of underlying, less gigantic, but nonetheless significant and perhaps more regular wave-like merger activity.

As a straightforward way to investigate this possibility, this section presents

	priors		posteriors				ML
	mean	std	mean	median	std	95%-band	MLE
μ_1	6.4	1000.00	5.804	6.274	1.179	(2.467, 6.451)	6.3129
μ_2	6.4	1000.00	6.604	6.588	0.206	(6.323, 7.037)	6.7007
σ_1	0.1	0.14	0.125	0.128	0.036	(0.053, 0.200)	0.1222
σ_2	0.1	0.14	0.105	0.105	0.030	(0.055, 0.161)	0.0753
ϕ_1	0.0	1.00	0.704	0.728	0.136	(0.399, 0.932)	0.4655
p_{11}	0.9	0.21	0.736	0.890	0.292	(0.078, 1.000)	0.9495
p_{22}	0.9	0.21	0.793	0.907	0.252	(0.115, 1.000)	0.8518

Note: Result of 20,000 Gibbs-Sampling iterations, iterations 1,000 through 20,000 used for inference. 95%-band refers to 95% posterior probability bands.

Table 2: Estimation Results for US Merger Activity, 1973:I through 1995:III.

estimates for the two-state model using only the data *prior* to the estimated break date which started the tidal wave, i.e. from 1973:I through 1995:III. Note that this corresponds to discarding little more than a quarter of the full series, which should leave us with sufficient data points to identify waves. Table 2 gives the parameter estimates resulting from fitting an AR(1) model. Comparison with estimation results from the full series in Table 1 shows that parameter inference is rather imprecise for the subsample. Figure 5 again shows the inferred probabilities of being in a high activity state for this particular subsample period. By comparison with the clear-cut result presented in Figure 3 for the entire sample period, Figure 5 suggests that waves are rather hard to identify in the merger data up to 1995:III. Although the plot does indeed hint at a somewhat increased probability of a high activity state around the mid 1980s (as well as around the mid 1990s as a warm-up to the ensuing large wave), this hint remains very faint due to the fact that, except for a short period around 1987, the probability of a high activity state stays in a rather tight band around 0.5. Overall, the fact that the inferred probability of a high activity state, $\Pr(S_t = 2|\mathcal{Y}_T)$, stays far clear of either 0 or 1 implies that the data reaches no clear conclusion concerning a likely sequence of states.²³

²³An alternative approach would be to extend our two-state model by introducing a third state, thereby explicitly allowing for both ‘medium’ and ‘high’ waves. However, casual inspection of the series strongly suggests that in such a three-state model, all quarters following 1995:III would rather clearly be associated with the ‘high wave state’, leading to inference on ‘low’ vs. ‘medium’ state comparable to the two-state analysis presented in this section.

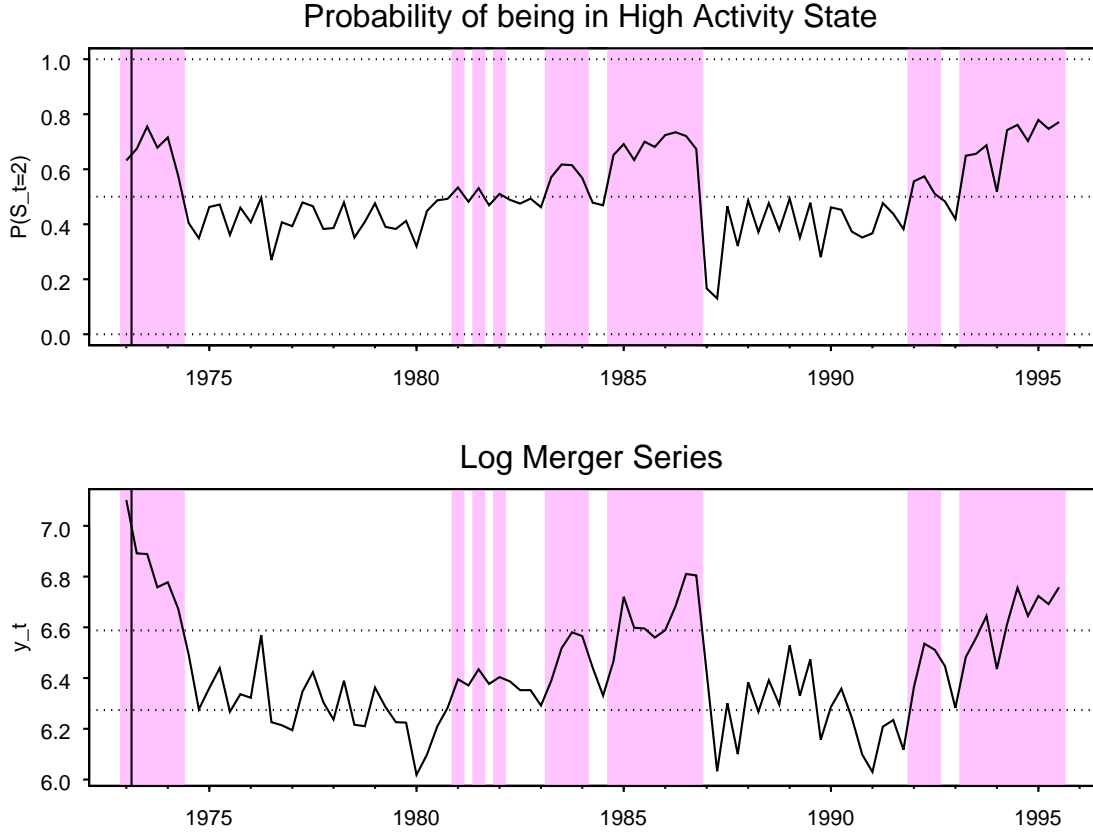


Figure 5: Estimated Probability of Being in State of High Merger Activity, US Log Merger Series 1973:I–1995:III (in Top Panel, Bottom Panel displays Log Merger Series).

This indecision in state inference is a first clear indication that the US have not witnessed a change in the mean level of merger activity at all during the 1980s.²⁴ A second striking feature of the probability plot in Figure 5 is its strong qualitative similarity with the underlying time series of log mergers, reproduced in the bottom panel of Figure 5. Indeed, the high-activity estimate in the top panel comes quite close to representing a positive affine transformation of the log merger series in the bottom panel. This is something we should expect to see from fitting a regime switching model to a series generated by a process with no actual switch in regime.

²⁴Note that in finite samples, given the data \mathcal{Y}_T , $P(S_t = 2|\mathcal{Y}_T)$ will of course generally deviate from 0.5 for any t even if the data were indeed generated by a stationary autoregressive process with no change in regime (i.e., a process with $\mu_1 = \mu_2$ and $\sigma_1 = \sigma_2$) due to remaining posterior uncertainty about the model parameters. Sample runs of the Gibbs Sampler on simulated data involving no regime change (and parameter values similar to those inferred for US mergers) revealed pictures quite similar to Figure 5.

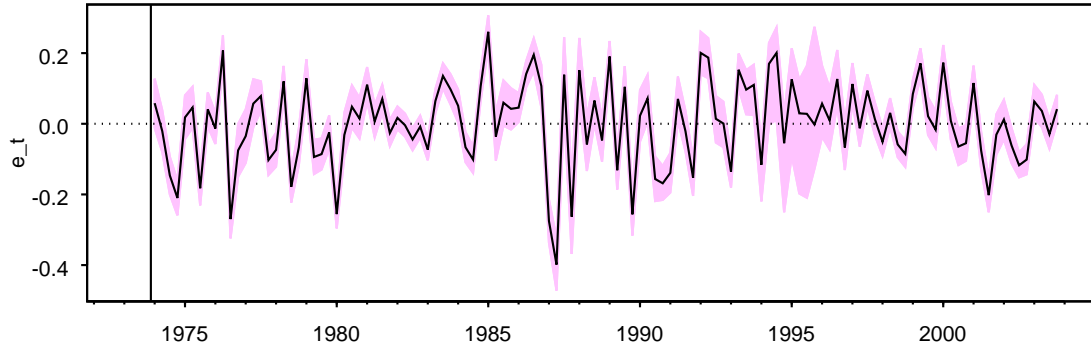


Figure 6: Mean Residuals (with 95% Posterior Bands) for US Log Merger Series 1973:I–2003:IV.

Third, Albert and Chib (1993) point out that fitting a Markov switching model to non-switching data results in rather large posterior bands on parameters, particularly concerning the means μ_1 and μ_2 and the transition probabilities p_{11} and p_{22} , which is reflected in our results. Finally and perhaps most importantly, inference runs with simulated data show the posterior distributions of $\mu_2 - \mu_1$ for our US subsample to be quite comparable to the posterior distributions of $\mu_2 - \mu_1$ resulting from simulated series with similar but constant model parameter values (i.e. a series with no state switch).

In sum, therefore, we find little evidence in support of the notion of an 1980s merger wave even in the truncated series. It is important however to be clear about the exact meaning of this result, as it does not contradict of course the 1980s having witnessed somewhat increased merger activity. What our result fleshes out is that this increased activity is rather unlikely to have been associated with a nonlinear shift in regime to the underlying autoregressive process (i.e. a ‘burst in activity’, such as the boom following 1995:III). Rather, it appears to be quite compatible with regular, well-behaved random shocks hitting a stationary linear autoregressive process. This is perhaps best illustrated by Figure 6, which plots the mean residuals (i.e. estimates of ε_t) resulting from the original estimation of the model over the full data range from 1973:I through 2003:IV.²⁵ Overall, these residuals appear well-behaved and compatible with our model assumptions of serial independence and

²⁵Recall that in a Bayesian estimation context, we consider posterior (i.e. ‘updated’) *distributions* of the parameters rather than particular point estimates. Thus, the resulting residuals themselves are random not only due to uncertainty about the unobservable state, but also due to posterior parameter uncertainty. Figure 6 displays both the mean and 95% posterior probability bands for the residual in any period.

normality. Nonetheless, one might indeed see the first half of the 1980s as having been hit by a sequence of slightly above-average shocks which—amplified by the processes’ strong positive autocorrelation—have given rise to a period of somewhat increased merger activity. We argue, however, that this is compatible with the usual behavior of a stationary autoregressive process rather than signifying a non-linear burst such as a discrete switch in mean over that period. In other words, in terms of our regime switching model of waves, we would like to suggest that the 1980s merger wave symbolized a ‘ripple’ rather than a real wave.

5.4 Parameter Uncertainty Matters: (Mis)identifying the 1980s Wave

The preceding section has thoroughly discussed our finding that, in contrast to the 1990s merger wave, the increased merger activity in the 1980s constituted no extraordinary burst in activity. This leaves unexplained, however, why the aforementioned studies by Town (1992) and Linn and Zhu (1997), both of which similarly fit a two-state mean-switching Markov model to US merger data, have identified an 1980s merger wave nonetheless.

Candidate reasons for this difference are manifold, as the studies differ in the particular series used, the time span considered, and details in model specification. As this section argues, however, the main difference in our interpretation of the 1980s merger wave is likely to stem from a more subtle, methodological reason: the refined methods of inference offered by the Gibbs sampling approach.

Recall that by means of this approach, our inference on regimes as presented in Figures 3 and 5 is based on $\Pr(S_t = 2|\mathcal{Y}_T)$. This contrasts with Town’s (1992) and Linn and Zhu’s (1997) inference based essentially on $\Pr(S_t = 2|\mathcal{Y}_T; \hat{\beta})$, the likelihood of a high state of merger activity while treating the underlying model parameters as given by the result of a preceding maximum likelihood estimation. Put differently, previous studies have asked the following question: Given that we believe the obtained parameter estimates to correspond to the true parameter values, what is the likelihood of a high activity state in any period? What this question obviously forgets to ask is just how reliable these parameter estimates actually are. Thus, if parameter estimates involve a high degree of uncertainty, answers to this question can severely overstate the evidence in favor of a high state of activity, and

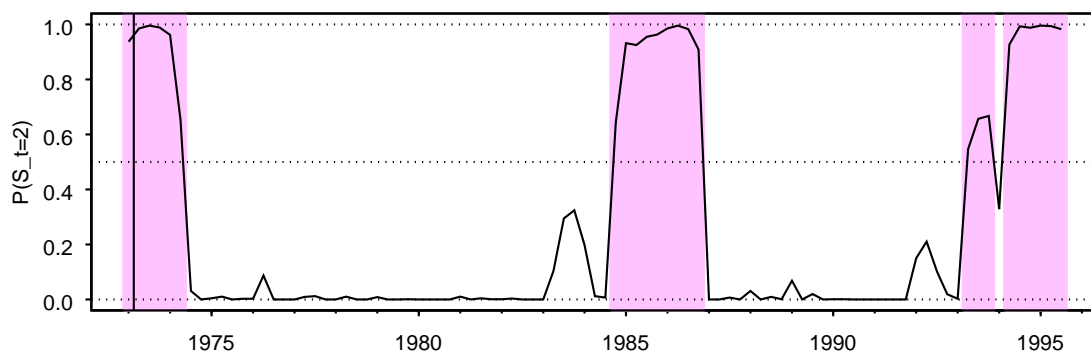


Figure 7: Probability of Being in High State of Merger Activity, US Log Merger Series 1973:I–1995:III, Calculated at Maximum Likelihood Parameter Estimates.

this appears to be highly relevant to the discussion of the 1980s merger wave.

To drive this point home, using our data spanning 1973:I through 1995:III, we have replicated the procedure in Town (1992) and Linn and Zhu (1997) by calculating the probability of a high activity state while holding the model parameters fixed at their maximum likelihood estimates reported in Table 2. Figure 7 plots the results. Clearly, by comparison with Figure 5, neglecting parameter uncertainty both leads to considerably more clear-cut inference on regimes and accentuates evidence for a high state of activity in the 1980s. Interestingly, the resulting sharp identification of a merger wave lasting from late 1984 to late 1986 is not very far from findings in Town (1992), who identifies 1986:IV as a period of intense merger activity, and findings in Linn and Zhu (1997), who identify the ‘mid-to-late 1980s’ as a merger wave.

5.5 But Waves Do Exist: The UK Merger Wave Experience

Next, we inspect the UK merger series, shown in the bottom panel of Figure 1, for its wave-like behavior. Analyzing the UK series turns out to be rewarding not only from the point of view of understanding UK merger activity, but also as a more general validation of our proposed Markov switching merger model and the wave hypothesis in particular. Indeed, the preceding analysis of the US merger data may be seen as somewhat disappointing as regards the wave hypothesis: While the proposed model itself does seem to provide a very good description of the US data, it does so in a fashion that hardly reflects the *repeated* bursts of activity attributed

	priors		posteriors				ML
	mean	std	mean	median	std	95%-band	MLE
μ_1	5.00	1000.00	4.789	4.823	0.342	(4.655, 4.957)	4.8328
μ_2	5.00	1000.00	5.756	5.781	0.325	(5.057, 6.037)	5.8324
σ_1	0.15	0.21	0.193	0.193	0.020	(0.141, 0.225)	0.1915
σ_2	0.15	0.21	0.173	0.169	0.033	(0.126, 0.255)	0.1497
ϕ_1	0.00	1.00	0.603	0.583	0.105	(0.455, 0.896)	0.5511
p_{11}	0.90	0.21	0.973	0.985	0.067	(0.892, 0.999)	0.9810
p_{22}	0.90	0.21	0.908	0.928	0.094	(0.724, 0.991)	0.9061

Note: Result of 20,000 Gibbs-Sampling iterations, iterations 1,000 through 20,000 used for inference. 95%-band refers to 95% posterior probability bands.

Table 3: Estimation Results for UK Merger Activity, 1969:I through 2003:IV.

to mergers by the literature—namely by identifying only a single switch from low to high activity around the mid 1990s, but no subsequent reversion to low activity, let alone a second or even a third wave. Given the limitation of our analysis to the last 30 years of US merger activity, this finding is of course not an invalidation of the wave hypothesis *per se*. However, the idea of a wave-like process governing US mergers would certainly be reinforced by observing more complete and possibly repeated wave cycles in other series such as the UK’s.

To analyze the UK data, we estimate an AR(1)-version of the model in Section 3. Table 3 reports summary statistics on both the priors used and on the estimated posterior distributions of the parameters. As with the full US series, the UK series permits a significant update on the model parameters’ priors, as shown by the marginal posteriors’ standard deviation and posterior bands. Again, the inferred high activity mean μ_2 significantly exceeds the low activity mean μ_1 . Interestingly, the ratio of high to low activity (in log terms) seems comparable across the US and the UK, as the high activity mean μ_2 exceeds the low activity mean μ_1 by an average (and mean) 20% for both series. Also, shocks to UK mergers again appear to be less volatile in the high activity state, although this is much less significantly so than for the full US series (as would be expected, fitting the model to the untransformed data reveals much higher volatility in the high than in the low activity state). The estimate of ϕ_1 suggests that autocorrelation in the UK series is significantly positive and again quite comparable to the degree observed in the US data. Finally, however,

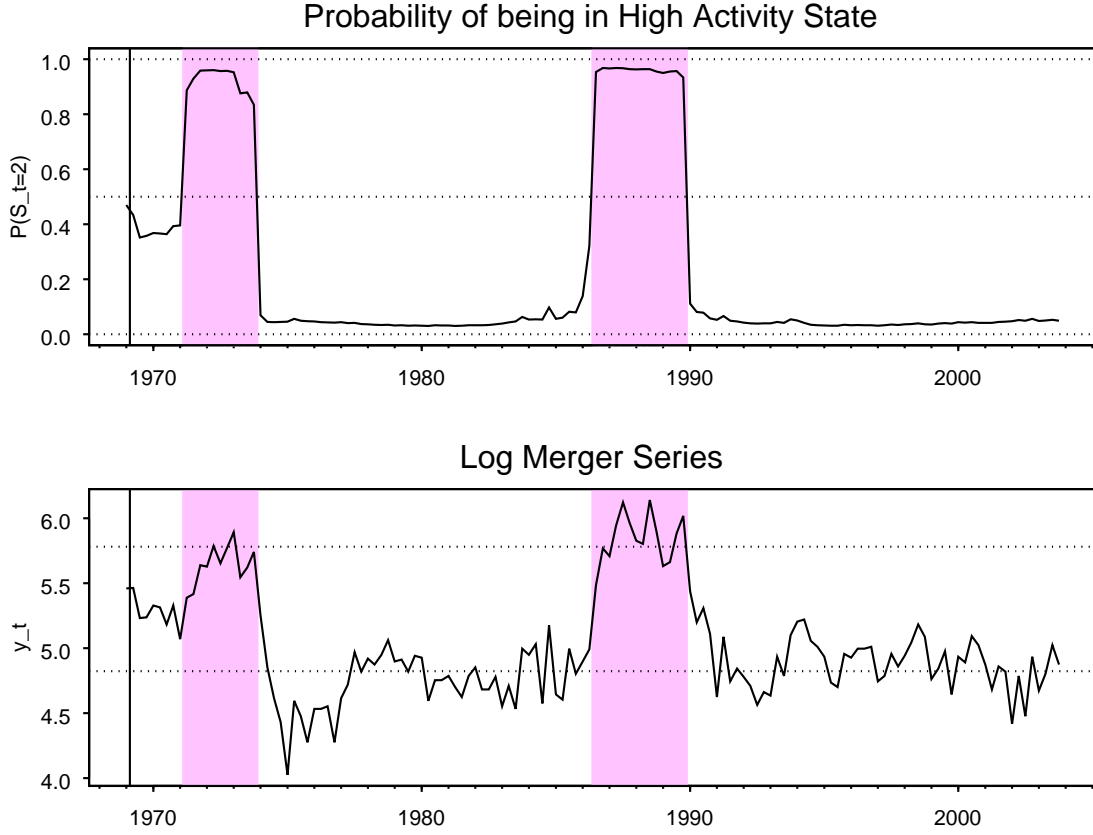


Figure 8: Estimated Probability of Being in State of High Merger Activity, UK Log Merger Series 1969:I–2003:IV (in Top Panel, Bottom Panel displays Log Merger Series).

the unobserved state is more likely to change in any period for the UK series, as shown by the posteriors on p_{11} and p_{22} . Particularly, the median expected duration of a low activity state is 16 years whereas the median expected duration of a high activity state is 13 quarters in the UK. Concerning the ergodic probability of being in a high state (unconditional on the data), $\Pr(S_t = 2)$, the mean posterior is 22.9%, whereas the median is 17.0%.

Next, Figure 8 shows the estimated probabilities of being in a high activity state for the UK merger series. The probability plot shows strong evidence that the UK has witnessed two distinct merger waves between 1969 and 2003. The first wave seems to have lasted from 1971:I through 1973:IV. Due to the aforementioned inference problems at the beginning of the series as well as because the merger series seems somewhat ‘undecided’ prior to 1971, we cannot precisely date the beginning of the first wave. Indeed, our inference leaves open to some extent whether the UK

started off in a high or low activity state in 1969.²⁶ It appears quite clear, however, that this first wave found its end in 1973:IV, as $\Pr(S_t = 2|\tilde{y}_T)$ drops from 0.836 in 1973:IV to 0.066 in 1874:I. The second wave in turn is reliably estimated to have started in 1986:III and ended in 1989:IV (the probability of a high state jumps from 0.320 in 1986:II to 0.955 in 1986:III and dips from 0.936 in 1989:IV to 0.109 in 1990:I).

6 Conclusion

The goal of this paper has been to revisit quantitative evidence on the merger wave hypothesis. Using a model of Markovian parameter switching, recent merger data and refined methods of inference, we have sought to identify and date waves in merger activity. A key finding has been that, concerning merger activity in the US and the UK over the past 30 years, the interpretation of merger activity as a mean and variance switching autoregressive process provides a fruitful quantitative operationalization of the wave hypothesis.

Moreover, fitting such a model to the data has produced the following merger wave chronology: First, since the beginning of our US series in 1973, the US appear to have witnessed only the beginning of a single large merger wave, this wave having been kicked off between the third and fourth quarter of 1995. Particularly, as a second major result and in contrast to other recent empirical work, we find very little evidence for the much discussed 1980s merger wave. We have argued that there is a methodological reason for this discrepancy in findings, as less refined inference methods which neglect parameter uncertainty are likely to have played a significant role in the econometric misidentification of 1980s merger activity as a wave. Third, fitting our model to UK merger activity between 1969 and 2003 clearly identifies two UK merger waves, one in the early 70s and a second in the late 1980s.

We hope that these findings will serve as a sound basis for a further discussion and investigation of possible underlying causes for merger waves. Particularly, the rather precise identification and precise timing of distinct states of merger activity based on our Markov-switching model openly calls for an economic interpretation and explanation of these states. Investigating one such hypothesis, our brief digression

²⁶In fact, by investigating annual data, Hughes (1993, p. 17) argues that merger activity culminates in twin peaks of activity in the late 1960s and early 1970s.

in Section 5.2 concerning industry-level data for the US has argued that, whatever the trigger for the US 1990s wave, industries seem to have been rather uniformly affected by it. Resende (1999), using an industry-level Markov switching model, reaches a similar conclusion concerning UK merger activity. Interestingly, while this leads us to conclude that nationally, industries in each the US and the UK seem to have been similarly affected by the observed waves, comparison of our results in Sections 5.1 and 5.5 quickly reveals that there is no sign of a similar ‘coordination of waves’ across countries over the last 30 years.

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