Growth, Savings, Financial Markets and Markov Switching Regimes

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Abstract: We report evidence that the relation between the financial sector share, private savings and growth in the United States 1948–1996 is characterized by several regime shifts. The finding is based on vector autoregressions on quarterly data that allow for Markov switching regimes. The evidence may be interpreted as support for a hypothesis that the relation between financial development and growth evolves in a stepwise fashion. Theoretical models where financial market extensions entail fixed costs imply such stepwise patterns. The estimated variable relations are roughly consistent with the patterns to be expected from such models, although our data do not admit definite conclusions. The timing of the shifts coincides with regulatory changes and changes in the financial market structure.

Keywords: Financial development, growth, Markov switching, savings, vector autoregression.

JEL Classification Numbers: C32, E44, O16, O51.

1. Introduction

There are theoretical and empirical reasons to expect that financial markets are extended stepwise. In theory, such market extensions would be characterized by a different relation to growth and saving. In practice, extensions are likely to take some time. Empirically we would then expect switches between a normal intermediate regime and a transition regime. Markov switching regressions are, therefore, natural tools to study whether such switches may have occurred. In this paper, we indeed find evidence supporting the regime switching hypothesis for quarterly U.S. times series 1948–1996.

Financial development can influence growth in three distinct ways: by raising the proportion of saving actually invested; by raising the social marginal productivity; by influencing the private saving rate. The first mechanism depends on the efficiency of financial intermediation, i.e., the fraction of saving absorbed to pay for financial intermediation services.

Date: This version: September 18, 1998.

Thanks: Thomas Lindh gratefully acknowledges financial support from *The Swedish Council for Research in the Humanities and Social Sciences (HSFR)*. Anders Warne gratefully acknowledges financial support from *Jan Wallanders och Tom Hedelius' stiftelse för samhällsvetenskaplig forskning*. Comments from participants at the 6th Annual Symposium of The Society for Nonlinear Dynamics and Econometrics, 1998, were very helpful for the final version.

The second mechanism works by improving the allocation of capital through information pooling like in Greenwood and Jovanovic (1990) or risk pooling, e.g. liquidity risks as in Bencivenga and Smith (1991). Financial risk pooling can also substitute for technological diversification like in Saint-Paul (1992) where financial markets allow an increased degree of specialization in production.

The third mechanism has ambiguous effects on growth since saving may go either way. If consumers have utility functions with a positive third derivative, precautionary saving decreases. Financial markets may also ease liquidity constraints on consumers by providing consumer credit that further reduces saving. On the other hand, financial development may increase the rate of return on saving and, thus, boosts saving.

Devereux and Smith (1994) and Obstfeld (1994) study models incorporating these mechanisms. Pagano (1993) is a succinct summary of research around these issues. Levine (1997) provides a more comprehensive survey of the literature on growth and financial development.

If there are inherent fixed costs in financial market extensions — as is very likely since they entail new information networks — such extensions will take place discretely. Saint-Paul (1992) and Lindh (1994) study such models. Alternatively, indivisibilities in the production technology may require a threshold level of financial development in order to be insurable by financial diversification, see Acemoglu and Zilibotti (1997). We would then expect to observe in the empirical record that sudden rises in the financial cost share are associated with sudden changes in the relation to growth and saving. The details in the resulting time series pattern will depend on the precise interaction of different mechanisms as well as other concurrent changes in the economy.

It is an empirical issue to study first whether regime shifts actually are present and, second, if any differences in the relation can be observed. This study accomplishes the first issue and gives some indications for the second. To these ends, we estimate Markov switching vector autoregressions using quarterly data 1948-1996 on changes in the financial sector share of U.S. corporate GDP, the growth rate of non-financial corporate GDP and the gross private saving rate.

There is evidence of several regime switches in a vector autoregressive model fitted to these data and the estimated relations in the intermediate regimes are largely consistent with theoretical expectations. Moreover, the shifts coincide in time with major changes in legislation and financial market structure in the U.S., adding credibility to the hypothesis that the shifts are not statistical artifacts, but reflect real economic regime changes.

The next section discusses in some detail the different patterns that could theoretically be expected in the relation between saving, growth and the financial sector share. Section 3 contains the empirical analysis. Section 4 examines the estimated regime process in more detail by reporting evidence from the financial market evolution in the U.S. Finally, Section 5 offers our conclusions.

2. Financial Markets and Growth

It is a stylized fact that financial development goes hand in hand with economic development. In the recent surge of cross-country regressions of growth on just about every conceivable variable, the positive correlation of financial development with growth is one of the few findings that seems reasonably robust to the inclusion of alternative sets of control variables.[1]

Causality is, however, still a matter of debate.[2] Lindh and Lindström (1997) report evidence that shifts in the financial sector share are associated with changing relations to growth and saving. Studying causality without due consideration of such regime shifts is liable to confound conclusions.

We first discuss the possible data patterns that could arise from financial regime shifts as an aid in the interpretation of the statistical model we use below to explore the nexus of financial development, growth and saving. We have no strong preferences for any specific model explaining the connection between financial development and growth. The discussion below is therefore heuristic and does not rely on a formal model. However, we do rely rather heavily on general features that characterize many recently studied models.

2.1. *Interpretative Framework*

The main point we wish to explore is the possibility that financial development proceeds by a sequence of shifts between transition regimes as financial markets are extended and intermediate regimes as the new market configurations are consolidated. Especially relevant in the previous literature are Saint-Paul (1992) and Acemoglu and Zilibotti (1997). They have developed theoretical models intended to explain a one-time shift from a stage of underdeveloped financial markets with highly variable production. Outcomes

[1]See Levine and Renelt (1992) generally about fragility analysis and Levine and Zervos (1993) on robustness of financial development measures.

[2]Jung (1986) as Demetriades and Hussein (1996) find causality to be mainly bidirectional. King and Levine (1993a, 1993b) conclude that there is a long run causality from financial markets to growth. However, Arestis and Demetriades (1997) argue that the cross-section evidence presented in the King and Levine papers is insufficient for causality analysis. Kugler and Neusser (1994) find long-run causality in time series from the financial sector to manufacturing TFP.

are then diversified by using less risky but on average less productive technology. Saint-Paul's results hinge on fixed information costs for establishing a financial market, such that it pays to switch to a more specialized technology only when enough capital has been accumulated. Acemoglu and Zilibotti, in contrast, assume that there are indivisibilities in the technology itself that force a slow rate of accumulation before the financial system is sufficiently deepened to insure the greater risks associated with high fixed investment costs.

The latter model is more elaborate, but both studies rely on production that is linear in capital — i.e. an endogenous growth model — and a non-convexity which introduces a threshold of capital accumulation below which high growth projects will not be undertaken because financial markets are unable to provide the necessary diversification of the higher risks associated with these projects. While Saint-Paul is based on a capital externality and a binary technology choice to generate endogenous growth, Acemoglu and Zilibotti assume that there is a pecuniary externality due to missing financial markets. In Saint-Paul's model there is, therefore, a direct link between the rate of saving and the rate of growth that necessitates an assumption of a very low risk preference parameter (below unity) in order to guarantee increased saving rates from financial development.

In Acemoglu and Zilibotti a heterogeneous set of risky projects carries the growth potential as more high-yield projects become feasible with more developed financial markets. Therefore the saving rate can be kept constant in the model by assuming logarithmic utility of consumption and still the model yields a similar result.

Generalizing Saint-Paul's basic model, Lindh (1994) points out that there may well be a sequence of fixed costs associated with extensions of the financial markets. That makes the idea relevant not only to the question of a one-shot growth takeoff, but also to variations in the relation between financial development, growth and savings in developed economies. Moreover, allowing for risk preferences that imply precautionary saving, growth enhancing effects may well be counteracted by saving declines. Empirical estimates of risk preference parameters are generally well above unity, hence, confirming precautionary saving.

Our approach is to study the relations between the financial sector share, φ, the growth rate g, and the rate of saving β. We will investigate whether data accept the hypothesis that the joint evolution of these variables in U.S. times series data is characterized by regime switches. In order to interpret the results it is useful to first state what data patterns we would expect to observe.

2.2. *Stylized Predictions*

Our study is mainly exploratory and we do not formally test any specific model. Nevertheless, in order to generate some predictions (see Table 1) we assume:

1. Financial market extensions entail a cost that is fixed in relation to the production level.
2. In the long run growth is increased by financial development, but this effect is comparatively slow.

The financial sector share measures the cost of transactions relative to the level of production. With a fixed volume of transactions per unit of production, growth would tend to decrease the financial sector share. An increased saving rate should tend to increase the volume of transactions and thus the financial sector share. Although these conclusions could be modified by scale and scope economies in transactions, changes in transaction technology and so forth, a fair guess is that φ increases with β and decreases with g in the intermediate regime. In the transition regime the cost hike due to added fixed costs will dominate and make predictions about saving and growth effects difficult.

Theoretically the saving rate is ambiguously affected by both growth and the financial sector share, but since the ambiguity derives from offsetting income and substitution effects in both cases, savings should at least be affected in the same way by both variables. As financial costs rise the direct effect on the saving rate is an unambiguous decrease. In the intermediate regime offsetting factors may dominate this leakage effect, but that is less likely in the transition regime.

Growth can initially be hampered by an increasing cost share for financial transactions. On the other hand more developed financial markets can increase the rate of capital accumulation and the rate of technological change. However, these positive effects would take some time, hence they will not show up in the transition regime. An increasing saving rate would be expected to increase long-run growth either transitorily or permanently but might have adverse effects on demand in the short run, thus depressing growth rates in the short run. On a quarterly frequency the demand effect could well be dominant.

In spite of the qualifications — indicated by question marks in Table 1 — the above arguments yield some guidance about the expected patterns of a vector autoregression. The effect on savings of financial extensions and faster growth is expected to reveal whether precautionary saving is dominant or not in the intermediate regime. Transition periods will be rather short compared to intermediate regimes and will tend to be triggered by high preceding rates of growth. In the absence of a formal model of the economic dynamics these predictions are not well defined hypotheses although they are informed by general traits of formal models in the literature.

3. The Empirical Analysis

In this section we examine empirical evidence for the intertemporal relationships between the extensions of financial markets, savings, and growth. In order to shed light on the issue discussed in the introduction we will let a vector autoregressive model (VAR) specialize into a Markov Switching vector autoregressive model (MS-VAR), and hence allow for different regimes to characterize the evolution of financial markets, saving, and growth.[3]

A priori we would expect one relatively frequent regime and one considerably less frequent regime, if it turns out that a two regime VAR seems fit to describe the data. There is one serious implication of this. Due to the limited amount of data available, the curse of dimensionality will effectively restrict the number of parameters in the model. As a consequence we will in the following only consider models with one lag, i.e. compare a VAR(1) with an MS-VAR(1), although results for a VAR(4) are also presented for comparison. Furthermore, and for the same reason, the precision of the estimates associated with the less frequent regime will be low and hamper interpretability of that regime.

3.1. *U.S. Quarterly Data*

Data are available on a quarterly frequency from 1946 to 1996 in NIPA (National Income and Product Accounts, U.S. Department of Commerce, 1992, 1997). However, the latest revision we had available only extends back to 1959 (taken from EconData's April 1997 update). Data from he period 1946-1958 are taken from U.S. Department of Commerce (1992).

We took the most reliable indicator of financial costs to be the gross domestic product attributed to financial corporate business.[4] It is not obvious how the relative cost should be measured. In order to avoid problems of interpretation and to enhance comparability we have used the financial sector share measured as financial GDP divided by corporate business GDP (Table 1.16 row 18 divided by row 1) in current values. This avoids the

[3] As can be seen from (1) below, it is more appropriate to label a VAR as a special case of an MS-VAR, i.e. the case of a single regime model. However, the above formulation can be justified considering the long tradition of econometric VAR models in contrast to the recently introduced MS-VAR model. An example of the latter is Blix (1998) examining Swedish inflation in a trivariate, two-regime model.

[4] The finance and insurance sector includes a number of real estate and business services that are not strictly financial. To sort out this we would need considerably more detailed industry divisions than are published in NIPA and most likely we would run into trouble with numerous changes in definitions over such a long period.

tricky issue of how government production should be treated in this context as well as ameliorating the problem of linking data between the revision and earlier series. The changes in definition as compared with earlier data mainly concern the government sector.

The saving share has for similar reasons been measured as gross private saving divided by the sum of private consumption and private domestic investment (Table 5.1 row 2 divided by Table 1.1 row 2 plus row 6). While the earlier NIPA convention essentially only added the budget surplus to gross private savings to arrive at gross savings the current convention adds actual gross government saving which makes up the bulk of the difference between the GDP measures. This raises the gross saving share quite considerably. Our measure is designed to avoid this problem. Similarly for comparability the growth rate has been computed as the growth rate of real non-financial corporate business GDP ($g_t = \log(y_t/y_{t-1})$ where y is taken from Table 1.16 row 36).

1959:4 and forwards real values are in terms of chained 1992 dollars. Before 1959:4 real values are in fixed 1987 dollars. The reason that the first three quarters in 1959 of the growth series are computed from U.S. Department of Commerce (1992) is that the first two quarters are missing in the 1997 revision. Furthermore the growth series starts only in 1948 since real value estimates are lacking for the first two years.

3.2. *The Statistical Model*

In order to make data suitable for the proposed analysis we have made the following transformations: the financial sector share and the national saving rate are in first differences, and hence stationary.[5] Moreover, by multiplying these two series by 400 and the growth rate by 100 we avoid numerical convergence problems in the estimations, and also get comparable measurement units, namely annual percentage points.

Let x_t be a trivariate time series with components $x_t = (\Delta\varphi_t,\ \Delta\beta_t,\ g_t)$, where $\Delta\varphi_t$ is the annual change in the financial sector share of GDP, $\Delta\beta_t$ is the annual change in the private saving rate, and g_t the annual growth rate of real non-financial corporate business

[5]Karlsen (1990, Chapter 5) gives a sufficient condition for second order stationarity which applies to MS-VAR models; see also Holst, Lindgren, Holst, and Thuvesholmen (1994). As long as the autoregressive coefficients depend on the regime process, the stationarity condition for linear VAR models is not valid and, hence, the existing unit root tests may not be meaningful. Moreover, there is not any theoretical guidance on how to relate the idea of cointegration to MS-VAR models under this circumstance. Since the growth rate looks stationary (see Figure 3), while the log of the financial sector share seems to be trending (Figure 1), the saving rate appears highly persistent (Figure 2), and hence that it is unlikely that the latter two series are cointegrated, we decided to apply first differences to the two possibly nonstationary time series. The first differences of these series are depicted in Figures 4 and 5, respectively. When we calculate the modulus of the largest eigenvalue, as defined by Karlsen, for the estimated first differenced MS-VAR(1) we find that it is roughly 0.43 and, hence, well inside the stationary region.

GDP. The vector x_t is assumed to be generated according to the following MS-VAR(p) model:

$$x_t = \mu_{s_t} + \sum_{k=1}^{p} A_{s_t}^{(k)} x_{t-k} + \varepsilon_t, \qquad t = 1, 2, \ldots, T, \tag{1}$$

where p is finite and typically small, $\varepsilon_t | s_t \sim N(0, \Omega_{s_t})$ with Ω_{s_t} being positive definite, and the initial values, $x_0, \ldots, x_{1-p}$, are taken as fixed.

The unobserved regime or state variable s_t is assumed to follow a q-state Markov process with transition probabilities $\Pr[s_t = j | s_{t-1} = i] = p_{ij}$, for all t and $i, j = 1, 2, \ldots, q$, and $\sum_{j=1}^{q} p_{ij} = 1$ is satisfied for all i. In addition, we assume that the Markov process is irreducible (no absorbing states) and ergodic.

For this particular application the maintained hypothesis is that $q = 2$, i.e. two states or regimes are sufficient for a fair description of the x_t process. We will, however, compare the two-regime model with a traditional single regime VAR, i.e. the case of $q = 1$. As noted above, we will, due to the small sample size, focus on models with one lag.

The random vector μ_{s_t} and the random matrices $A_{s_t}^{(k)}$ and Ω_{s_t} depend only on the state taken on by s_t. If $s_t = 1$, then $\mu_{s_t} = \mu_1$, $A_{s_t}^{(k)} = A_1^{(k)}$ and $\Omega_{s_t} = \Omega_1$. Maximum Likelihood (ML) estimates for the MS-VAR(1) model are obtained via the EM algorithm; for more details the reader is referred to Hamilton (1990, 1994). Standard errors for the point estimates are based on conditional scores, as in Hamilton (1996). The VAR(1) and the VAR(4) models are estimated with (Gaussian) ML.

Due to the presence of unidentified nuisance parameters under the null (the transition probabilities p_{ij} and the parameters of, say, the second regime) it is, as of yet, not clear how to test the single regime model against the two-regime model.[6] However, it is still possible to empirically discriminate between the single regime VAR models and the MS-VAR(1) by examining their performances in terms of specification tests, e.g. test for serial correlation and autoregressive conditional heteroskedasticity.

3.3. *Specification Results*

Table 2 presents some stylized facts about the behavior of the change in the financial sector share of GDP, the change in the saving rate, and the growth rate. Over the sample period we find that the change of the financial sector share of GDP each quarter is 0.1 percent with a standard deviation of about 0.8 percent. The average change per quarter in the saving rate is approximately zero, but this series is considerably more volatile than

[6]Some procedures have been suggested in the literature, for instance Hansen (1992).

the other two; it has an estimated standard deviation of roughly 4 percent. The average growth rate is much higher, about 1.75 percent, with a standard deviation of about 2.8 percent. Moreover, the three variables do not seem to be contemporaneously correlated as indicated by the small covariances.

Having estimated the MS-VAR(1) model, we may calculate the corresponding state conditional moments; see Warne (1996a) for details on the relationship between the state contingent moments and the parameters of the MS-VAR model. First, we find that about 168 observations belong to Regime 1 and the remaining 26 to Regime 2. The probability to remain in State 1 is estimated to be 95 percent and the corresponding estimate for State 2 is 66 percent. The point estimates of the Markov transition probabilities indicate that both regimes are persistent and that the second regime occurs less frequently than the first. An F-statistic version of the Wald test of the hypothesis that the Markov chain is not serially correlated, i.e. that the transition probabilities are equal to the long run or ergodic probabilities, is strongly rejected by the data when inference is based on the assumption of asymptotic normality.

Table 3 displays the conditional first two moments for x_t. The most apparent difference between the two states is the greater volatility recorded for the less frequently occuring second state, standard deviations are roughly 3 times larger. As for first moments, it can be seen that the mean of g_t is slightly larger in the stable first state, whereas the mean of $\Delta\beta_t$ is negative in the first state and in the second, volatile state, positive and considerably larger than the unconditional mean. Finally, the second state is associated with a decline in the financial sector share. The two states can be described as one stable and one volatile where saving increases and the financial sector share decreases.

Prior to an interpretation of the estimated state conditional parameters of model (1) it is useful to evaluate the data describing properties of the MS-VAR(1) model. We will use the VAR(1) as a reference model and examine how the two models conform to the assumptions of serially uncorrelated residuals and no autoregressive heteroskedasticity. Test results are summarized in Table 4.

According to the univariate specification tests the VAR(1) model is severely misspecified with respect to serial correlation, ARCH, and normality. This is not surprising when looking at the differenced data in Figures 3–5 which displays periods, or clusters, of high volatility. The null hypotheses for the MS-VAR(1) model can only be rejected in the case of ARCH in the financial sector share equation. Moreover, when testing the VAR(1) as a system, the multivariate tests suggest rejection of multivariate normality and serial correlation (p-values are .000 and .001 respectively). A multivariate ARCH test in the MS-VAR(1) model cannot be rejected, nor a multivariate test of serial correlation (p-values are .239 and .691, respectively).

Based on these tests our conclusion is therefore that the MS-VAR(1) provides an adequate description of the data, whereas the VAR(1) does not. In order to check that this result is not simply an effect of a larger set of parameters in the MS-VAR(1), or from serial correlation in the VAR(1), we undertook the same tests for a VAR(4). The results show that this larger model is still mis-specified, and the periodic volatility outbursts are not accounted for even with four lags, while there is no evidence of serial correlation.

3.4. *The Theoretical Predictions Meet the Data*

The results in Table 5 indicate effects in the intermediate regime (Regime 1) that are in rather good agreement with the theoretically expected pattern in Table 1. The financial sector share is positively affected by previous savings and negatively by previous growth as expected. The saving share is negatively (although insignificantly) affected by both the other variables. Although the sign could not be a priori determined the negative effect is in fact the one expected from empirical work on the elasticity of intertemporal substitution indicating that income effects dominate the saving response. Growth is related negatively to previous changes in the financial sector share and positively to the saving rate.

The estimated coefficients in the transition regime (Regime 2) are also close to the expected. However, being very imprecisely estimated, due to the few "observations" of that regime, we cannot attach much importance to the point estimates. A glance at Figure 1 indicates that the estimated transition regimes are catching downturns *after* a financial sector expansion rather than the expansions themselves.

One implication for the MS-VAR of the hypothesized pattern of effects in Table 1 is that the financial sector share should Granger cause the growth rate in mean-variance. Technically, this means that for some time periods

$$E\big[u_{g,t}^2\big|\{\Delta\varphi_\tau, \Delta\beta_\tau, g_\tau\}_{\tau=1-p}^{t-1}\big] \neq E\big[\tilde{u}_{g,t}^2\big|\{\Delta\beta_\tau, g_\tau\}_{\tau=1-p}^{t-1}\big],$$

where $u_{g,t} = g_t - E\big[g_t\big|\{\Delta\varphi_\tau, \Delta\beta_\tau, g_\tau\}_{\tau=1-p}^{t-1}\big]$, and $\tilde{u}_{g,t} = g_t - E\big[g_t\big|\{\Delta\beta_\tau, g_\tau\}_{\tau=1-p}^{t-1}\big]$; Warne (1996b) presents the set of necessary and sufficient conditions for Granger noncausality in mean (the standard Granger noncausality hypothesis), mean-variance, and distribution (conditional independence). Similarly, the pattern of effects in Table 1 implies that the financial sector share should be Granger causal for the saving rate in mean-variance.[7]

For the MS-VAR model, there are two channels through which, say, $\Delta\varphi$ can be useful for predicting the g. First, it can help to predict the regime (an "indirect" prediction channel). Second, conditional on the regime, it can help improving the one-step ahead forecast of the growth rate. Since there are two channels through which the financial sector share can be informative about the next period value (and the uncertainty of the prediction error), there is not a unique set of parameter restrictions for testing the noncausality hypothesis. However, there is a finite number of cases and if one of these cases is true, then $\Delta\varphi$ is Granger noncausal in mean-variance for the variable we are interested in. Given an MS-VAR model with 2 regimes and 3 observable variables, the total number of such cases is four.[8]

In the case of, say, the hypothesis $\Delta\varphi \not\Rightarrow g$, a common feature of the four sufficient conditions is that, conditional on the regime and the past values of $\Delta\beta$ and g, $\Delta\varphi$ does not help to predict the next period value of g. Letting a_{ij,s_t} denote the (i,j):th element of $A_{s_t}^{(1)}$, this means that $a_{31,s_t} = 0$ for both regimes. These restrictions represent the second prediction channel that was mentioned in the previous paragraph.

In Table 6 we report F-statistics and p-values from testing the hypothesis that $a_{ij,s_t} = 0$ for both values of s_t and with $i,j = 1,2,3$. The evidence here agrees with the results in Table 5 and, in particular, we cannot reject the hypotheses that the coefficients on the lagged financial sector share are zero for both the saving rate and the growth rate equations.

Next, in Table 7 we report F-statistics of Granger noncausality in the six different directions that are possible in our model. For a given pair, e.g. the financial sector share and the saving rate ($\Delta\varphi \not\Rightarrow \Delta\beta$), the noncausality hypothesis implies that at least one of

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[7]The predictions in Table 1 do not, however, imply that the financial sector share should be Granger causal for the growth rate or the saving rate in *mean*, i.e. the variances of $u_{g,t}$ and of $\tilde{u}_{g,t}$ can be equal. The reason is that if the Markov process is serially uncorrelated, then Granger noncausality in mean implies that the expected value of the random coefficient on $\Delta\varphi_{t-1}$ in the saving rate and in the growth rate equation, respectively, are zero. As long as the two possible values for each random coefficient have opposite signs, the weighted (by the ergodic probabilities for the Markov process) sum of the two values can be zero for each case. This is consistent with the hypothesized relationships in Table 1. Noncausality in mean-variance, however, requires that each possible value for these random coefficients is zero, and is therefore not consistent with the pattern of effects in Table 1.

[8]Under the assumption of conditional normality for ε_t in equation (1) and that the matrix with Markov transition probabilities has either full rank or rank equal to one (which is always satisfied when there are two regimes), Warne (1996b) shows that Granger noncausality in mean-variance is equivalent to Granger noncausality in distribution.

the four sets of restrictions, (C1.1), (C1.2), (C2), and (C3), must be satisfied; the specific parameter restrictions are given in the Table.

For the case when we wish to test the hypothesis that $\Delta\varphi$ is Granger noncausal in mean-variance for $\Delta\beta$, the restrictions (C1.1), (C1.2), and (C2) imply that $\Delta\varphi$ is conditionally uninformative about the regime, while (C3) implies that $\Delta\beta$ does not directly depend on the regime (other than via g or the residual covariances). Moreover, the (C1.1) and (C1.2) restrictions allow the Markov process to be serially correlated, while (C2) does not. Finally, (C1.1) and (C1.2) are different in the sense that (C1.1) means that *only* $\Delta\varphi$ has to be conditionally uninformative about the Markov process, while (C1.2) implies that both $\Delta\varphi$ and g are conditionally uninformative about the regime process. The former case turns out to imply that $\Delta\varphi$ is Granger noncausal in mean-variance for g as well as for $\Delta\beta$, while the latter case happens to imply that both $\Delta\varphi$ and g are Granger noncausal in mean-variance for $\Delta\beta$.[9]

From Table 7 we find that the (C1.1), (C1.2), and (C2) restrictions are strongly rejected by the data for both the saving rate and for the growth rate case. This suggests that the financial sector share process contains unique information for predicting the regime process. Exactly how it matters depends on which restriction(s) is (are) not consistent with the data. However, at the 5 percent level of marginal significance, we can only reject the (C3) hypothesis in the case of the saving rate. In other words, the financial sector share seems to be Granger causal (in mean-variance) for the saving rate, but not for the growth rate. Moreover, the growth rate appears to be Granger causal for the saving rate.

Apart from Granger causality from $\Delta\varphi$ and from g to $\Delta\beta$, there is no evidence of such causal links. Interestingly enough, it is always the (C3) restrictions that cannot be rejected for these cases, while the other sets of restrictions are always rejected. This means that all three variables are useful for making inference about the regime process and that the evidence on causality is primarily found in the equation describing the saving rate. The volatility outbursts, evident in Figures 3–5, can be found in all three time series and, thus, explain the former result.

4. What Happened at the Transitions?

Do the estimated transition regimes have any connection to real events? To answer that question we have surveyed parts of the literature on the development of the U.S. financial system, and organized that information around the five transition regimes with duration

[9] Notice that these two sets of restrictions are *not* nested.

greater than two quarters. As can be seen in Figure 6 there is a spike in the second quarter of 1958 that could be caused by the fact that we have had to link data around that point. The switch in the beginning of the 1990s is followed by two further indications of regime shifts in the second halves of 1994 and 1995. The latter of these is considerably less apparent than other switches. On the other hand the estimated two instances of the transition regime in the beginning of the 1980s are probably linked. Below, we consider them to be one transition period.

The U.S. financial system became increasingly regulated in the 1930s and during the war. The McFadden Act 1927 attempted to disallow interstate banking and the Glass-Steagall Act 1933 the combination of investment and commercial banking. Moreover, state legislation in many cases prohibited even branching within states. Deposit interest rates were bounded from above until quite recently, although the restriction did not become binding until the late 1960s. During the first decades after World War II these regulations were upheld and loopholes were plugged (for example the Douglas Amendment to the Bank Holding Act in 1956 prohibits bank holding companies from owning banks in more than one state). In the end of the 1960s the regulation framework started to erode gradually. Internationally much of the postwar experience is dominated by the buildup of the Bretton-Woods monetary system and its subsequent breakdown.

The interaction between financial developments and legislation is, of course, a mutual interdependence. Changes in the financial system necessitate changes in legislation, in turn precipitating new changes in the financial system. However, the causal connection is by most observers seen as running from financial system pressures to legislation. Deregulation has to a large extent been motivated by practices that already had started to evolve within the bounds of the old rules.

Below we very briefly try to specify some events which we believe are connected to switches in financial regime.

4.1. *The Transition Periods*

1950:1–51:3. Consumer credit controls were abolished in June 1949 and interest ceilings on deposits were successively raised through 1948. In June 1950 the Korea War began, leading to a short speculative boom. Credit outstanding from the Federal Reserve increased very fast in 1949-1950 (and also in 1958-59, see Friedman & Schwartz, 1963, Chart 54). In 1950 federal insurance of savings and loans associations was raised to the same level as that for commercial banks. Although at a very modest scale, charge account banking started to develop in 1950, see Klebaner (1990, p. 203). Record mortgage volumes were accompanied by liberalization of government mortgage programs in 1950, and savings bank statutes were amended to allow out-of-state federally underwritten mortgages (Klaman, 1961).

1975:1–75:4. The state of Maine allowed out of state bank holding companies in 1975. When negotiated brokerage commissions were allowed in 1975 the structure of investment banking rapidly changed. In 1973 the final breakdown of the Bretton-Woods system made new services for international payments necessary. In 1974 automated brokerage services were made available to primary dealers in Treasury securities. Trade in interest rate futures were introduced in Chicago 1975, see Wilson (1993).

The first automated clearing houses were established in 1972. During the period 1971–73 regulation of negotiable certificates of deposits were abolished leading to massive volume increases in the trading of these instruments up to 1975. 1972–73 NOW-accounts started to evolve, i.e. saving accounts with negotiable orders of withdrawal, in effect an interest-bearing checking account. Regular demand deposits had a zero interest rate ceiling under Regulation Q.

Local government (November 1974) and business corporation (November 1975) savings deposits and telephone transfers to cover checks (April 1975) were allowed. This was all part of a general movement to circumvent interest regulations on demand deposits. In 1972 the first mutual money market funds also appeared. See Klebaner (1990) for more details.

1981:3–82:3. In December 1980 shelf registration of bond issues was allowed, i.e. anticipatory clearing of bond issues. The market for so called junk bonds (lacking normal credit rating) started to increase rapidly, from a volume of around 1.2 billion dollars in 1981 to 30.9 billion dollars in 1989. In 1980 interest rate regulation started to phase out over the following six years. Money market deposits were allowed in 1982. Automated teller machines made branching regulation less efficient. Mortgage-backed securities and federally backed variable rate mortgages were allowed. Market instability in 1982-84 forced the Fed to insure all deposits rather than only small ones. 1980–82 savings and loans associations were given more discretion in the range of services they could offer, for details see Meerschwam (1991) and Mullineux (1987).

In 1981 International Banking Facilities were exempted from reserve requirements and from some state and local government taxation in order for U.S. banks to be competitive in international markets. A Supreme Court ruling in 1981 accelerated merger activities as regulators were forced to become more liberal. In 1982 the Garn-St Germain Act allowed thrifts to offer banking services previously reserved for commercial banks. At the other extreme investment banks offered cash management accounts competing with commercial banks. There was a general tendency of dissolving the distinctions between local thrifts, mainly savings and loan associations, commercial banks and investment banks. This coincided with an increased failure rate that slowed down the pace of deregulation. Mullineux (1987) discusses this in detail.

In 1982 Congress responded to market pressures and doubled the statutory limits on bankers acceptances to 200 percent of bank capital. Towards the end of 1982 money market deposit accounts were introduced. Securitization of international debt as well as loan sales, interest swaps and other innovative financial activities expanded, see Klebaner (1990). Securitization, i.e., the packaging and trading of debt on second-hand markets had previously (since the 1960s) mainly been applied to federally guaranteed home mortgages. Options on a diverse array of futures contracts were introduced at several exchange markets in 1982.

1992:2–93:1 & 1994:2–3. After a 1985 precedent in the United States Supreme Court regional banking was allowed and it expanded rapidly due to state deregulation at the end of the decade, see Hawawini and Swary (1990). As a response to the problems of thrift institutions, mainly savings and loan associations in the 1980s, the previous regulatory authorities were dissolved in 1989 and replaced by the Office of Thrift Supervision which in the following years reduced the reliance on deposit funding that had caused much of the problems, see Stiroh (1997).

1991 the federal deposit insurance was reformed towards risk based insurance. 1992 world-wide futures and options trading was introduced. In 1993 the junk bond market went through a revival. 1994 the Riegle-Neal Interstate Banking and Banking Efficiency Act finally imposed federal rules allowing bank holding companies and commercial banks to establish branch units across state lines. More details can be found in OECD (1992-1995).

4.2. *Summary of Evidence*

Every regime shift is associated with new developments on the financial markets and the regulatory framework. However, the shifts are also associated with upturns from recessions in economic activity. Upturns from recessions though are not always associated with regime shifts. The Markov transitions in the data mark increased volatility in the recessions. This could be interpreted to mean that we simply pick out recession phenomena, due to other causes than financial development. But as the brief review above shows, the transitory regimes are coinciding with or succeeding important changes in the financial markets and their regulation.

It is not inconsistent with theory to assume that new extensions of the financial markets are associated with a period of increased volatility and recession as the market settles down to a new competitive equilibrium. The pattern found in the transitory regime, even if statistically not well determined, still points in the direction that growth and saving recovers in the transition before the switch back to the intermediary regime. At the same time, inflated financial costs at the beginning of the transition tend to subside to a lower and perhaps more sustainable level. In that interpretation the first switch to the transition regime is triggered by financial costs that outrun the level that can be sustained by saving and growth in the economy. The transition phase is then an equilibrating recovery phase which again triggers a switch back to a calmer development phase.

5. Conclusions

The main conclusion from our work is that the MS-VAR model successfully describes the data since it can account for much of the heteroskedasticity that is apparent in the data. This indicates that the possibility of regime switches should be taken into account when analyzing the time series relations between financial development and growth.

The autoregressive patterns in the intermediate regimes are largely consistent with theoretical expectations. These patterns indicate that the main effect of financial development is to decrease precautionary savings. The patterns in the transition regime — although not well determined — are widely conforming to the expected.

There is plenty of institutional evidence that the statistically determined transition periods are associated with new developments on the financial markets on several levels. Substantive changes in regulation as well as rapid developments in new financial instruments and markets are closely associated with the timing of the transition regimes.

Further research along these lines may shed new light on the old question of causality between financial development and growth. Our Granger causality analysis points in the direction that causal (predictive) effects can only be discovered from the financial sector share and from growth to the saving rate. This is consistent with the finding that all three variables are useful when attempting to make inferences about the regime process.

It should be emphasized that the noncausality hypothesis concerns only one quarter ahead predictions and that our results have little (if anything) to say about longer forecast horizons. Since the length of a time period in the theoretical discussion is unspecified (although it is not far fetched to take it to be longer than one quarter), the lack of Granger causality from the financial sector share to the growth rate should not be interpreted as strong evidence against the hypothesis of a causal link from financial development to growth. It suggests, however, that the dynamics of the link need to be better specified in order to settle the issue.

TABLE 1: Hypothesized pattern of effects in different regimes.

	intermediate			transition		
	φ_{t-1}	β_{t-1}	g_{t-1}	φ_{t-1}	β_{t-1}	g_{t-1}
φ_t	?	+?	$-$?	?	?	?
β_t	+/$-$	?	+/$-$	$-$	?	$-$?
g_t	+	+($-$)?	?	$-$	$-$	?

TABLE 2: Estimated state independent moments.

variable	mean	covariances		
$\Delta\varphi_t$	0.103	0.693	-0.612	-0.784
$\Delta\beta_t$	0.020	-	15.91	0.917
g_t	1.758	-	-	8.005

TABLE 3: Estimated state dependent moments for a 2-state MS-VAR(1) model.

	Regime 1			
variable	mean	covariances		
$\Delta\varphi_t$	0.145	0.316	-0.342	-0.655
$\Delta\beta_t$	-0.058	-	7.527	0.098
g_t	1.805	-	-	5.960
	Regime 2			
variable	mean	covariances		
$\Delta\varphi_t$	-0.165	3.007	-2.169	-1.695
$\Delta\beta_t$	0.518	-	68.84	6.283
g_t	1.457	-	-	20.88
	The Markov Process			
	$\hat{p}_{11} = .947$		$\hat{p}_{22} = .663$	
	(.032)		(.157)	
	$\hat{\pi}_1 = .864$		$\hat{T}_1 = 167.7$	
	(.063)		(12.2)	
H_0:	$p_{11} + p_{22} = 1$		$F = 39.68$	
			[.000]	

TABLE 4: Univariate specification tests for a MS-VAR(1), a VAR(1), and a VAR(4) with respect to serially uncorrelated residuals, normality, and no autoregressive heteroskedasticity

	serial correlation		normality		ARCH	
VAR(1)	statistic[a]	p-value	statistic[b]	p-value	statistic[c]	p-value
$\Delta\varphi_t$	2.36	.042	105.17	.000	3.79	.006
$\Delta\beta_t$	2.93	.014	32.66	.000	16.69	.000
g_t	1.18	.321	20.64	.000	5.79	.000
VAR(4)	statistic[d]	p-value	statistic[b]	p-value	statistic[e]	p-value
$\Delta\varphi_t$	1.31	.263	100.39	.000	2.40	.052
$\Delta\beta_t$	1.38	.234	11.41	.003	13.78	.000
g_t	0.49	.783	25.44	.000	2.04	.092
MS-VAR(1)	statistic[f]	p-value	statistic	p-value	statistic[g]	p-value
$\Delta\varphi_t$	1.23	.298	-	-	2.98	.020
$\Delta\beta_t$	1.46	.216	-	-	1.29	.275
g_t	1.03	.396	-	-	2.26	.064

a is an F(5,185)-statistic for an LM test of serial residual correlation as reported by Pc-Fiml 9.0

b is a $\chi^2(2)$–statistic for normality based on excess skewness and kurtosis as reported by Pc-Fiml 9.0

c is an F(4,182)-statistic for an LM test of ARCH based on lagged squared residuals as in Pc-Fiml 9.0

d is an F(5,173)-statistic for an LM test of serial residual correlation as reported by Pc-Fiml 9.0

e is an F(4,170)-statistic for an LM test of ARCH based on lagged squared residuals as in Pc-Fiml 9.0

f is an F(4,178)-statistic for a test of serial residual correlation based on conditional scores

g is an F(4,178)-statistic for a test of ARCH based on conditional scores

TABLE 5: ML Estimates of μ_{s_t}, $A_{s_t}^{(1)}$, $s_t = 1, 2$, for the MS-VAR(1) model; estimated standard errors based on conditional scores within parentheses, and significant coefficients in bold.

equation	μ_1	$A_1^{(1)}$			μ_2	$A_2^{(1)}$		
$\Delta\varphi_t$	**0.148**	**0.236**	0.010	-0.018	-0.060	-0.612	0.012	-0.093
	(0.071)	(0.077)	(0.017)	(0.024)	(0.740)	(0.713)	(0.241)	(0.352)
$\Delta\beta_t$	0.117	-0.340	-0.004	-0.073	2.161	0.416	-0.676	**-0.888**
	(0.338)	(0.384)	(0.057)	(0.119)	(2.572)	(2.225)	(0.444)	(0.408)
g_t	**1.332**	-0.062	0.010	**0.269**	0.866	-0.525	-0.173	0.391
	(0.295)	(0.361)	(0.069)	(0.083)	(1.720)	(3.737)	(0.275)	(0.818)

TABLE 6: F-tests of the hypothesis that $a_{ij,s_t}^{(1)} = 0$ for both regimes with p-values within brackets.

equation	variable $\Delta\varphi_{t-1}$	$\Delta\beta_{t-1}$	g_{t-1}
$\Delta\varphi_t$	4.597	.189	.316
	[.011]	[.828]	[.730]
$\Delta\beta_t$	.370	1.095	2.452
	[.691]	[.337]	[.089]
g_t	.026	.187	5.261
	[.975]	[.830]	[.006]

The reference distribution is $F(2, T-12)$ and the F-statistic is computed as $F = ((T-12)/(2T))W$, where W is the Wald statistic.

TABLE 7: F-tests of various Granger noncausality in mean-variance hypotheses with p-values in brackets.

		$\Delta\varphi \nRightarrow \Delta\beta$	$\Delta\varphi \nRightarrow g$	$\Delta\beta \nRightarrow \Delta\varphi$	$\Delta\beta \nRightarrow g$	$g \nRightarrow \Delta\varphi$	$g \nRightarrow \Delta\beta$
H_0	r	F	F	F	F	F	F
(C1.1)	13	3.383	3.383	2.564	2.564	3.085	3.085
		[.000]	[.000]	[.003]	[.003]	[.000]	[.000]
(C1.2)	19	2.568	3.293	2.969	3.293	2.969	2.568
		[.001]	[.000]	[.000]	[.000]	[.000]	[.001]
(C2)	3	5.375	4.583	4.304	3.956	3.903	5.535
		[.001]	[.004]	[.006]	[.009]	[.010]	[.001]
(C3)	6	2.602	.604	1.131	.603	1.200	2.443
		[.019]	[.726]	[.346]	[.728]	[.308]	[.027]

The $F(r, T-s)$-approximated F-statistic is computed from $F = ((T-s)/(Tr))W$, where W is the Wald statistic, r is the number of restrictions, and s is the closest integer to the average number of free parameters per equation under H_0, i.e. $s = \text{int}[(38-r)/3]$.

For (C1.1) the reference distribution is $F(13, T-8)$. The set of restrictions is: $\mu_{i,1} = \mu_{i,2}$, $a_{i1,1} = a_{i1,2}$, $a_{i2,1} = a_{i2,2}$, $a_{i3,1} = a_{i3,2}$, $a_{ji,1} = 0$, $a_{ji,2} = 0$, $a_{ki,1} = 0$, $a_{ki,2} = 0$, $\omega_{ii,1} = \omega_{ii,2}$, $\omega_{ij,1} = 0$, $\omega_{ij,2} = 0$, $\omega_{ik,1} = 0$, $\omega_{ik,2} = 0$ for the hypothesis $x_i \nRightarrow x_j$, where $k \notin \{i, j\}$ while $x_1 = \Delta\varphi$, $x_2 = \Delta\beta$, and $x_3 = g$.

For (C1.2) the reference distribution is $F(19, T-6)$. The set of restrictions is: $\mu_{i,1} = \mu_{i,2}$, $\mu_{k,1} = \mu_{k,2}$, $a_{i1,1} = a_{i1,2}$, $a_{i2,1} = a_{i2,2}$, $a_{i3,1} = a_{i3,2}$, $a_{k1,1} = a_{k1,2}$, $a_{k2,1} = a_{k2,2}$, $a_{k3,1} = a_{k3,2}$, $a_{ji,1} = 0$, $a_{ji,2} = 0$, $a_{jk,1} = 0$, $a_{jk,2} = 0$, $\omega_{ii,1} = \omega_{ii,2}$, $\omega_{ik,1} = \omega_{ik,2}$, $\omega_{kk,1} = \omega_{kk,2}$, $\omega_{ij,1} = 0$, $\omega_{ij,2} = 0$, $\omega_{kj,1} = 0$, $\omega_{kj,2} = 0$ for the hypothesis $x_i \nRightarrow x_j$, where $k \notin \{i, j\}$.

For (C2) the reference distribution is $F(3, T-12)$. The set of restrictions is: $p_{11} = p_{21}$, $a_{ji,1} = 0$, $a_{ji,2} = 0$ for the hypothesis $x_i \nRightarrow x_j$.

For (C3) the reference distribution is $F(6, T-11)$. The set of restrictions is: $\mu_{j,1} = \mu_{j,2}$, $a_{ji,1} = 0$, $a_{ji,2} = 0$, $a_{jk,1} = a_{jk,2}$, $a_{jj,1} = a_{jj,2}$, $\omega_{jj,1} = \omega_{jj,2}$ for the hypothesis $x_i \nRightarrow x_j$, where $k \notin \{i, j\}$.

FIGURE 1: Financial sector share of corporate GDP with the estimated transition regime periods, $\Pr[s_t = 2|\mathcal{X}_T; \hat{\theta}] > 0.5$, in the shaded areas.

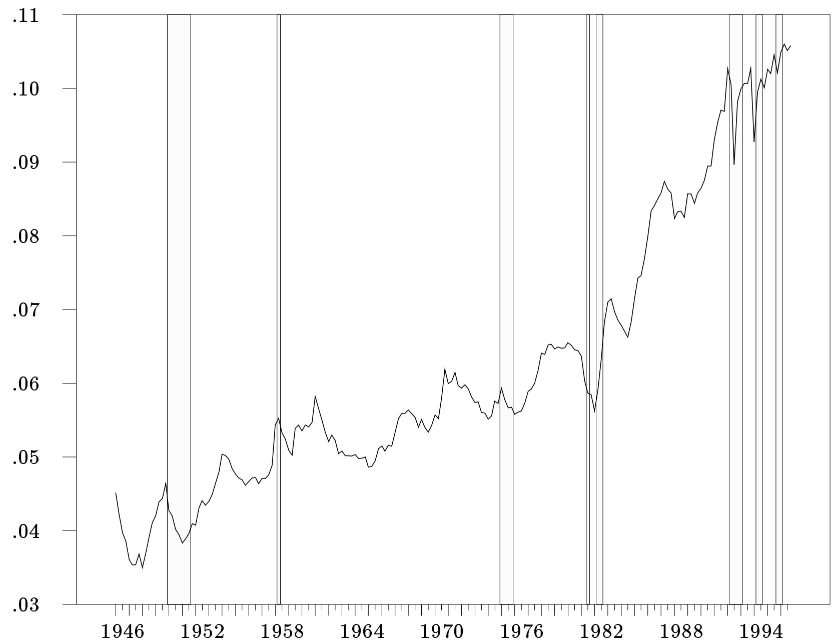


FIGURE 2: Gross private saving rate with the estimated transition regime periods, $\Pr[s_t = 2|\mathcal{X}_T; \hat{\theta}] > 0.5$, in the shaded areas.

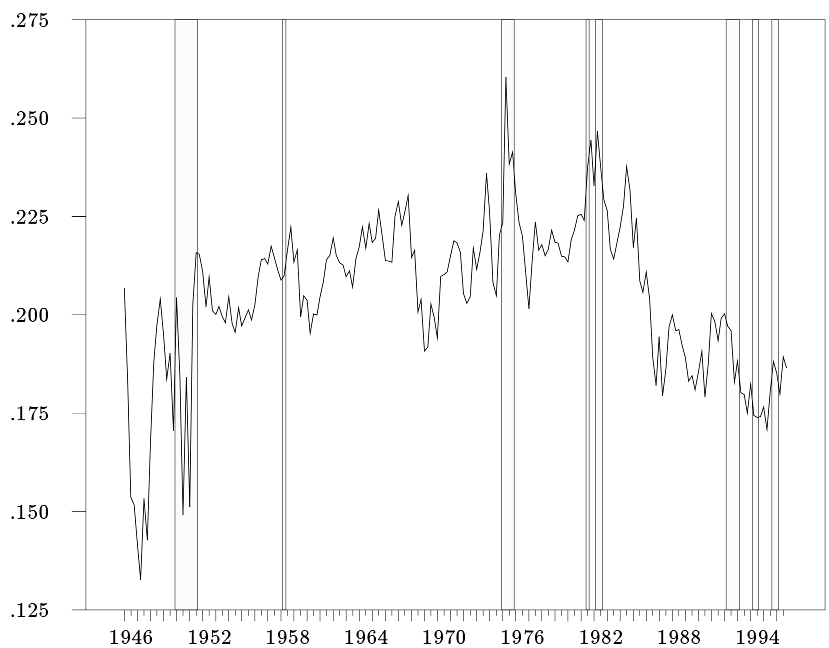


FIGURE 3: Growth rate of non-financial corporate GDP with the estimated transition regime periods, $\Pr[s_t = 2|\mathcal{X}_T; \hat{\theta}] > 0.5$, in the shaded areas.

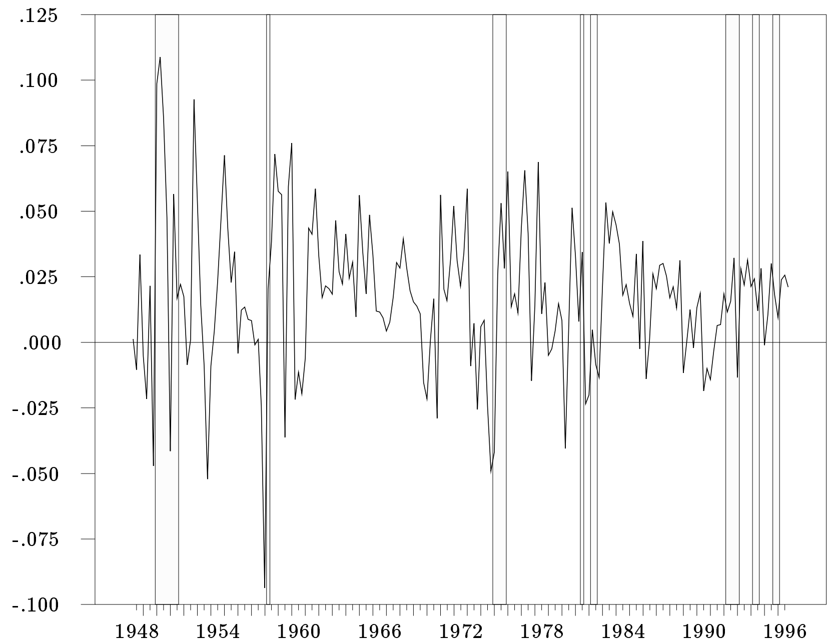


FIGURE 4: First differences of financial sector share of corporate GDP with the estimated transition regime periods, $\Pr[s_t = 2|\mathcal{X}_T; \hat{\theta}] > 0.5$, in the shaded areas.

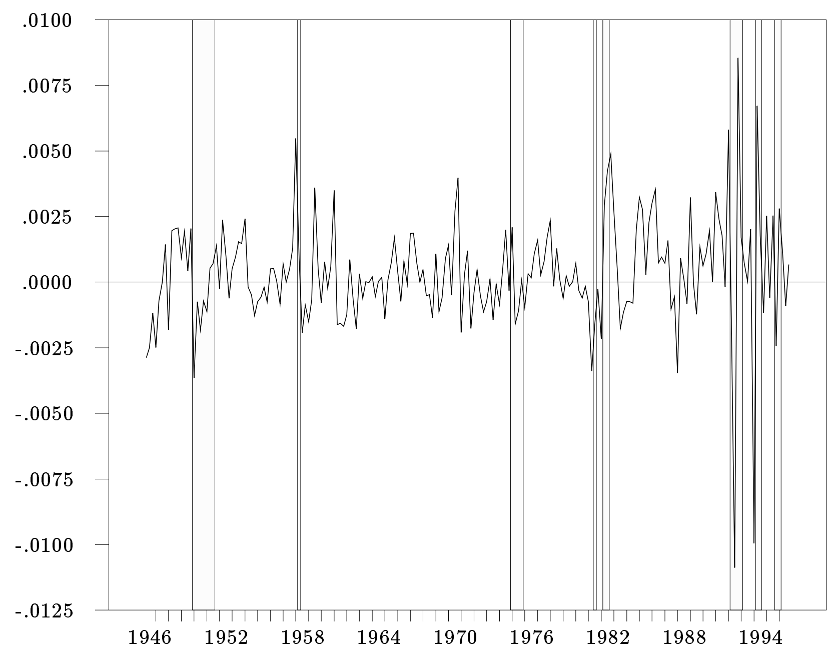


FIGURE 5: First differences of the gross private saving rate with the estimated transition regime periods, $\Pr[s_t = 2|\mathcal{X}_T; \hat{\theta}] > 0.5$, in the shaded areas.

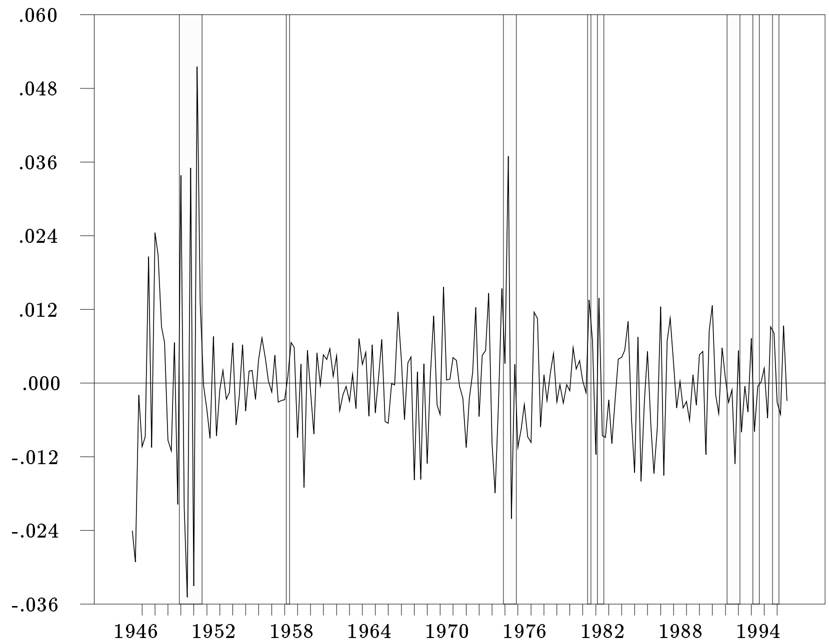
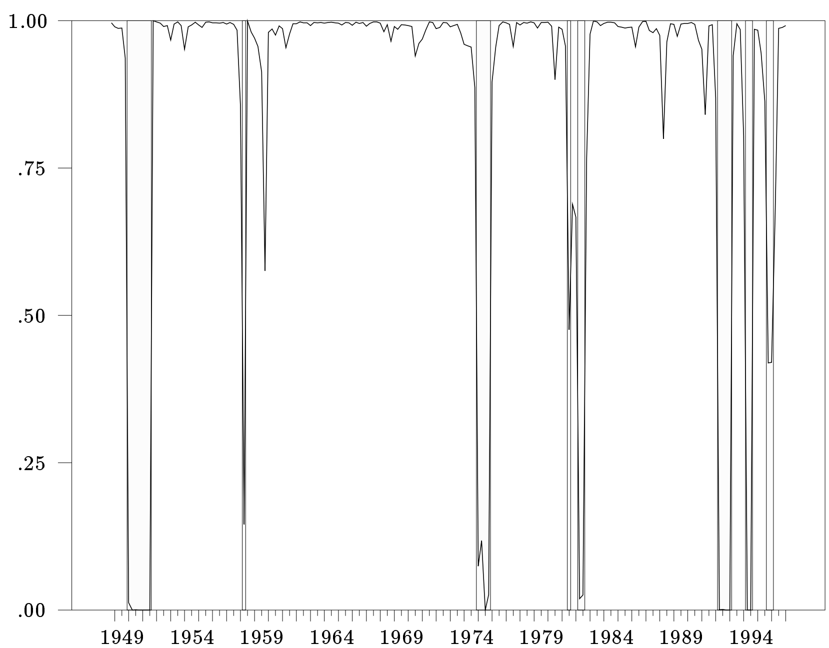


FIGURE 6: Estimated smooth probabilities of being in the intermediate regime with the estimated transition regime periods, $\Pr[s_t = 2|\mathcal{X}_T; \hat{\theta}] > 0.5$, in the shaded areas.



References

Acemoglu, D., and Zilibotti, F. (1997). "Was Prometheus Unbound by Chance? Risk, Diversification and Growth." *Journal of Political Economy*, *105*, 709-751.

Arestis, P., and Demetriades, P. (1997). "Financial Development and Economic Growth: Assessing the Evidence." *Economic Journal*, *107*, 783-799.

Bencivenga, V. R., and Smith, B. (1991). "Financial Intermediation and Endogenous Growth." *Review of Economic Studies*, *58*, 195-209.

Blix, M. (1998). *Forecasting Swedish Inflation with a Markov Switching VAR.* (Mimeo, Sveriges Riksbank)

Demetriades, P. O., and Hussein, K. A. (1996). "Does Financial Development Cause Economic Growth? Time Series Evidence from 16 Countries." *Journal of Development Economics*, *51*, 385-409.

Devereux, M. B., and Smith, G. W. (1994). "International Risk Sharing and Economic Growth." *International Economic Review*, *35*, 535-550.

Friedman, M., and Schwartz, A. J. (1963). *A Monetary History of the United States 1867-1960.* Princeton: Princeton University Press.

Greenwood, J., and Jovanovic, B. (1990). "Financial Development, Growth, and the Distribution of Income." *Journal of Political Economy*, *98*, 1076-1107.

Hamilton, J. (1990). "Analysis of Time Series Subject to Changes in Regime." *Journal of Econometrics*, *46*, 39-70.

Hamilton, J. (1994). *Time Series Analysis.* Princeton: Princeton University Press.

Hamilton, J. (1996). "Specification Testing in Markov Switching Time Series Models." *Journal of Econometrics*, *70*, 127-157.

Hansen, B. (1992). "The Likelihood Ratio Test under Nonstandard Conditions: Testing the Markov Switching Model of GDP." *Journal of Applied Econometrics*, *7*, S61-S82.

Hawawini, G., and Swary, I. (1990). *Mergers and Acquisitions in the U.S. Banking Industry: Evidence from the Capital Markets.* Amsterdam: North-Holland.

Holst, U., Lindgren, G., Holst, J., and Thuvesholmen, M. (1994). "Recursive Estimation in Switching Autoregressions with a Markov Regime." *Journal of Time Series Analaysis*, *15*, 489–506.

Jung, W. (1986). "Financial Development and Economic Growth: International Evidence." *Economic Development and Cultural Change*, *34*, 158-167.

Karlsen, H. A. (1990). *A Class of Non-Linear time series models.* PhD Thesis, Department of Mathematics, University of Bergen.

King, R., and Levine, R. (1993a). "Finance, Entrepreneurship, and Growth: Theory and Evidence." *Journal of Monetary Economics*, *32*, 513-542.

King, R., and Levine, R. (1993b). "Finance and Growth: Schumpeter Might Be Right." *Quarterly Journal of Economics*, *108*, 717-737.

Klaman, S. B. (1961). *The Postwar Residential Mortgage Market.* Princeton: Princeton University Press.

Klebaner, B. J. (1990). *American Commercial Banking: A History.* Boston: Twayne Publishers.

Kugler, M., and Neusser, K. (1994). *Manufacturing Growth and Financial Sector Development in OECD Countries.* (Working Paper 1/94, Europa-Universität Viadrina, Frankfurt (Oder), Fakultät für Wirtschaftswissenschaften)

Levine, R. (1997). "Financial Development and Economic Growth: Views and Agenda." *Journal of Economic Literature*, *35*, 688-726.

Levine, R., and Renelt, D. (1992). "A Sensitivity Analysis of Cross-Country Growth Regressions." *American Economic Review*, *82*, 942-963.

Levine, R., and Zervos, S. (1993). "What We Have Learned About Policy and Growth from Cross-Country Regressions?" *American Economic Review*, *83*, 426-430. (Papers and Proceedings)

Lindh, T. (1994). *Financial Market Costs, Growth, and the Trade-Off with Risk of Specialization.* (Uppsala University, Department of Economics, Working Paper 1994:31)

Lindh, T., and Lindström, T. (1997). "Structural Breaks in the Financial Sector Share, Savings and GDP Growth in OECD: 1960-1992." *Applied Economics*, *29*, 33-41.

Meerschwam, D. M. (1991). *Breaking Financial Boundaries: Global Capital, National Deregulation, and Financial Services Firms.* Boston: Harvard Business School Press.

Mullineux, A. (1987). *International Banking and Financial Systems: A Comparison.* London: Graham & Trotman.

Obstfeld, M. (1994). "Risk-Taking, Global Diversification, and Growth." *American Economic Review*, *84*, 1310-1329.

OECD. (1992-1995). *Financial Market Trends.* Paris: OECD.

Pagano, M. (1993). "Financial Markets and Growth: An Overview." *European Economic Review*, *37*, 613-622.

Saint-Paul, G. (1992). "Technological Choice, Financial Markets and Economic Development." *European Economic Review*, *36*, 763-781.

Stiroh, K. J. (1997). "The Evolution of an Industry: U.S. Thrifts in the 1990s." *Journal of Banking and Finance*, 1375-1394.

U.S. Department of Commerce. (1992). *National Income and Product Accounts of the United States, I.* (Bureau of Economic Analysis, Washington D.C., Data released January 1993).

U.S. Department of Commerce. (1997). *National Income and Product Accounts of the United States.* (Bureau of Economic Analysis, Washington D.C., Data released April 1997).

Warne, A. (1996a). *Autocovariance Functions and Maximum Likelihood in a VAR Model under Markov Switching.* (Manuscript, Institute for International Economic Studies, Stockholm University)

Warne, A. (1996b). *Causality in a Markov Switching VAR.* (Manuscript, Institute for International Economic Studies, Stockholm University)

Wilson, J. (1993). *Money Markets: The International Perspective.* London: Routledge.

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