How Democratic Transitions Affect the Saving-Investment Correlation: Evidence from Taiwan

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ABSTRACT

In this paper, we introduce a DCC GARCH-X model to study the impact of democratic transitions on the correlation between savings and investment in Taiwan. We find that democratic transition lowers the correlation through decreasing variances of saving and investment, and financial openness raises it through increasing the covariance counterpart. Our empirical findings suggest that democratic transitions are crucial to explaining why the correlation between savings and investment in developing countries is higher than that in developed countries.

1. Introduction

The purpose of this paper is to investigate how democratic transition affects the saving-investment correlation in Taiwan. Feldstein and Horioka (1980) used data in major industrial countries to measure the extent to which a higher national saving rate in a country is associated with a higher rate of domestic investment rate. They found that differences in national saving rate have corresponded to almost equal differences in domestic investment rates. The high saving retention ratio is not consistent with the implication of financial openness in these countries. On the other hand, Dooley et al. (1987), Wong (1990), Coakley et al. (1999) and Kasuga (2004) used data in developing countries to conduct similar studies and found that the saving retention ratios are significantly lower than those in developed countries. Since the degree of financial openness is low in most developing countries, financial openness alone cannot account for the differences in saving retention ratios between developing and developed countries. Given that developed countries are democratic countries, can democracies account for the difference?

Although the Taiwanese economic miracle has sometime been used as the case for arguing that authoritarian regimes can better pursue aggressive economic development, the peaceful democratic transition (the lift of martial law in 1987, the general parliament election in 1992, the first direct presidential election in 1996, the alternation of political forces in power in 2000) has been hailed as the most significant achievement in its post-war political economy development. In the meantime, we also witnessed the rising correlation between savings and investment. This raises the question of whether the link between democracy and the correlation between savings and investment exists. In this paper, I find that democratic transitions affect not only the

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long-run trends in national savings rate and domestic investment rate but also the correlation between these two variables through the channel of the variances of these two variables.

Almost all work on democracies and national economic performance has focused on the empirical link between democracy and long-run growth, with contradictory findings. Many channels of positive influence from democracy on growth have been proposed. First, by encouraging an open debate over the choice of public policies and policy makers, democracies discourage extremism and the take-over of power by illegitimate means. Hence, democratic institutions have the most transparent rules for the alternation of political forces in power among alternative political regimes. The lower degree of uncertainty associated with the peaceful and predictable transfer of political power is likely to foster investment and growth. Second, rulers with discretionary power tend to adopt distortional policies that benefit a small group of insiders at the expense of the general public. Democracies make it easier to keep these abuses in check and improve the quality of policy-making, by submitting politicians to regular public scrutiny and promoting viable alternatives in the form of opposition parties. Finally, by better securing property rights and facilitating contract enforcement, democracies may generate a higher rate of returns on physical investment and thus raise the incentives for private investment and growth.

The linkages between democracies and short-run economic performance have received much less attention in the literature. Sah (1991) argued that decentralized political regime (and the democratic institution in particular) should be less prone to volatility for the following reason. The presence of a wider range of decision makers results in a greater diversification and therefore less risk in an environment rife with imperfect information. Rodrik (1997) stressed that the democratic institution is the most effective mechanism of conflict management to reduce the effects of unavoidable external shocks on short-run economic performance in a society with deep social cleavages. Therefore, democracies induce macroeconomic stability. Recently, Quinn and Woolley (2001) proposed the voter’s preference hypothesis as follows. Assume that voters are risk-adverse. They tend to penalize incumbent governments for short-run volatility. In non-democracies, elite are more likely to seek risk that voters could reject. As a result, autocracies produce systematically more volatility than do democracies, which implies that democracies can reduce macroeconomic volatility, other things being equal. Empirical results indicate that voters have only modestly rewarded incumbent governments for higher economic growth rates but severely penalized them for increased economic volatility.

Even though co-movements among macroeconomic variables are the central focus of studies in aggregate fluctuations, the relationship between democracy and the co-movements receives very little research attention. Given the channels of positive influence of democracy on investment, the effect of democratic transitions on the correlation between savings and investment depends on how democracies affect the saving rate. If democracies are thought to spur a desire for current consumption (Huntington and Dominguez, 1975), democratic transitions may weaken the correlation through lowering the saving rate. On the other hand, in a democratic country, a strong middle class associated with a well functioning income redistribution system is likely to reduce the circumstances whereby the poor can expropriate the wealth of the rich. Democratic transitions can enhance the correlation through raising the saving rate.

To identify the channels of influence from democracy to the correlation between savings and investment, I use the democracy index, alongside of trade openness and financial openness, as the predetermined variables in Engle (2002)'s dynamic conditional correlation generalized autoregressive conditional heteroskedasticity model (here and henceforth, DCC GARCH-X model). Trade openness and financial openness are included in DCC GARCH-X model for the following reasons. First, the degree of trade openness can be influenced by the extent of political freedom. Protectionist policies tend to be imposed because they benefit a few firms at the expense of the general public. Democracies may weigh the general public preferences more heavily than autocracies, and therefore have less protectionism. The trade openness is included to investigate whether the link between democracies and the correlation is via this channel of influence. Second, the democracy could lower the entry barrier in industries (Aghion et al., 2007). For examole, in the democratic transition process, Taiwanese government eased controls on capital account, allowed the entry of new private banks in the domestic banking sector, and opened its door for foreign portfolio investment. If the democratic transition promotes financial openness, then the measurement of financial openness in the DCC GARCH-X model helps to clarify the causal relationship between financial openness and the correlation between savings and investment.

The plan for the remainder of the paper is as follows. Section 2 reviews empirical regularities on the long-run trend and short-run volatility of both national savings and domestic investment rates and the correlation between these two variables in the period 1971:1 to 2008:2. To remove the trend, the linear time trend, the index of democracy and the measurement of financial openness are used as regressors in regression equation. Regression results suggest that democracy is a significant determinant of the trend in the levels of national saving and domestic investment rates. Section 3 describes the specification of DCC GARCH-X model and the two-step estimation procedure. In Section 4, I present empirical results, which indicate that democratic transition lowers the correlation through decreasing variances of saving and investment, and financial openness raises it through increasing its covariance counterpart. The implications of the results are considered in a concluding section.

2. Democracy and the trends of saving and investment rates

In this section, I first present the evidence that the national savings rate and the domestic investment rate have different trends. Then, I propose a new trend-removing procedure, based on the concept of long-run trend in Kydland and Prescott (1990), in which linear time trend, the democracy index and other institution variables are independent variables in the trend-removing regression. Finally, I will show empirical regularities of the rolling statistics constructed from the residuals in the trend-removing regressions.

Figure 1: National saving rate, Domestic investment rate and the Balance on current account as shares of GDP

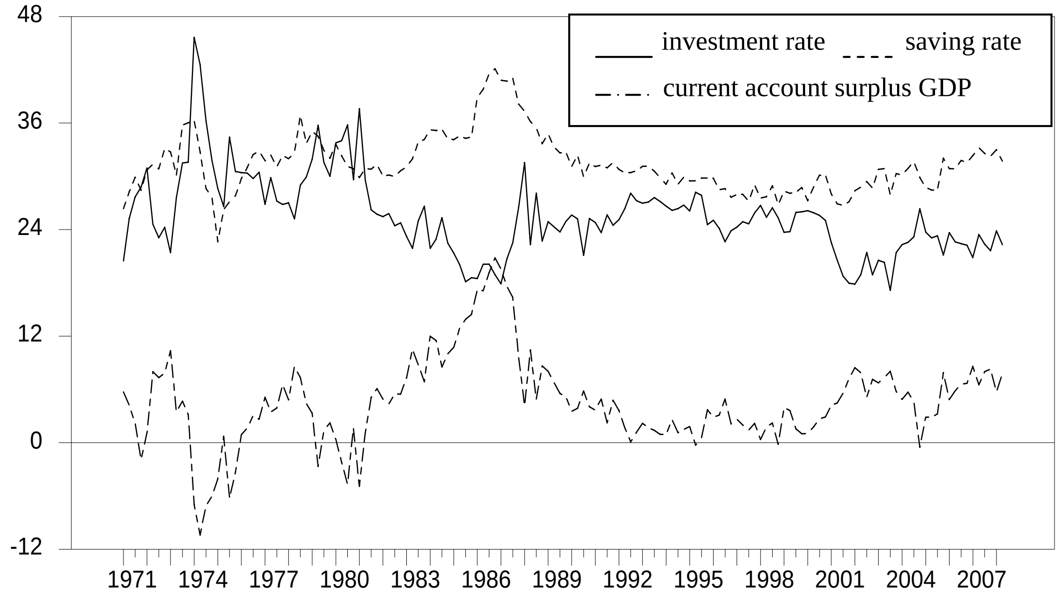


Figure 1 plots national saving rate, domestic investment rate and the balance on current account as shares of GDP for Taiwan over the period 1971:1 to 2008:2. Three empirical regularities are evident from the plots. First, there were not only considerable ranges in both the national saving rate and the domestic investment rate, but also significant variations in the difference between the two series. Before 1980, national saving and domestic investment rates moved closely with a narrow variation in the difference between them. After that, national saving rate was significanly higher than domestic investment rate with a wide gap between them, which indicated significant capital outflows in this period. Second, national saving rate moved more closely with the balance on current account as shares of GDP than with domestic investment rate. With the perfect capital mobility, a country can borrow in the international capital market to finance its domestic investment projects. Therefore, movements in national saving rate should be reflected in those in the balance on current account as shares of GDP, which implies a weak statistical relationship between savings and investment. In the absence of statistical discrepancy, the identity (national savings – domestic investment = the balance on current account) holds. It is clear from the identity that the larger volatility of the balance on current account also yields a weaker statistical relationship between national saving and domestic investment rates. As clearly revealed in Figure 1, the increased volatility in the balance on current account as shares of GDP accounted for a lower correlation between savings and investment during the latter part of the 1980s and the 1999-2008 periods. Third and last, both national saving and domestic investment rates exhibited different long-term trends. If the co-integration relationship does not exist between these two series, the specification of the regression equation given in Feldstein and Horioka (1980) is not appropriate for the time series data.

Next, I will conduct various stationarity and unit root tests for national saving rate, domestic investment rate and the balance on current account as shares of GDP to detect their trend properties. The null (alternative) hypothesis of stationarity (unit root) test is that the data series under test is trend stationary with a linear time trend. When the stationarity test fails to reject its null and the unit root tests reject their null, the data series under test is said to be a trend stationary process with a linear time trend. Otherwise, the data series under test is difference stationary. Test results in Table 1 determine that both the national saving rate and the balance on current account as shares of GDP after removing the linear time trend are difference-stationary, but render weaker support for the hypothesis that the domestic investment rate is trend-stationary with a linear time trend.

Table 1 : Stationarity and Unit Root Tests

Variable	KPSS stationarity test	ADF unit root test	DF-GLS unit root test	PP unit root test
national saving rate	0.27**	-2.72	-1.98	-3.39*
domestic investment rate	0.12	-4.08***	-2.78*	-5.14***
the balance on current account as shares of GDP	0.23***	-2.55	-2.52	-3.08
s_t	0.09	-3.66**	-3.23**	-3.98**
i_t	0.08	-4.65***	-3.39**	-5.97***
ca_t	0.10	-3.45**	-3.18**	-4.47***
co-integration test	0.08	-5.68***	-1.54	-5.78***

Note :

1. s_t, i_t and ca_t are residuals taken from the trend-removing regressions for national saving rate, domestic investment rate and the balance on current account as shares of GDP, respectively.
2. *,** and *** indicate the null is rejected at the 10%, 5% and 1% level of significance, respectively.

Lin (2000) developed a model based on the subsistence level of consumption to account for the upward trend in the private saving rate. In his framework private consumption must first exceed the subsistence level of consumption, letting the intertemporal consideration guides households' consumption decisions only for that portion of their incomes left after the subsistence has been satisfied. The implied intertemporal elasticity of substitution will be close to zero for countries at or near subsistence consumption levels, and will rise thereafter. The rising intertemporal elasticity of substitution induces a rising saving rate. As the economy matures, the importance of subsistence consumption decreases in the households' decisions; and the intertemporal elasticity of substitution becomes stationary.

Feldstein (1994) argued that companies make foreign direct investments (FDI) in response to a variety of direct business needs -- being close to customers, obtaining lower cost labor, responding to pressure from government where sales occur -- rather than as a way of shifting capital from countries with lower to countries with higher marginal products. He found that outbound FDI reduced domestic investment by an approximately equal amount based on 18 OECD countries for the decade of the 1980s. As revealed in the figure on Taiwan's outbound FDI and domestic investment in Appendix, there is also a reverse relationship between these two series. It is in outbound FDI toward the PRC that the domestic investment rate exhibits a downward trend since the 1990s.

According to the results of the stationarity and unit root tests reported in Table 1, it is reasonable to model both the savings and investment rates as being difference-stationary with drift. When economic time series are difference stationary with drift, these series can be decomposed into linear time trend and stochastic trend components, and the notions of stochastic and deterministic co-integration are useful in examining the long-run interactions between two or more variable of interest. I now focus on the long-run relationship between these variables. As displayed in the last low of Table 1, the test result fails to reject the stochastic co-integration relationship between the national saving and domestic investment rates are, but rejects the deterministic co-integration. Since the effects of democracy uoon the trend property of the national savings and domestic investment rates were not considered in those tests, a weaker long-run relationship is expected to exist between them once these effects are taken into account.

The above findings are consistent with those reported in the previous literature. For example, De Haan and Siermann (1994) found evidence against the co-integration relationship between the savings and investment rate in Taiwan based on the annual series in the whole sample period 1951 to 1989, but failed to reject the co-integration relationship for the period 1951 to 1980. It is consistent with the empirical regularity that national saving and domestic investment rates moved relatively closely with narrow variation in the difference between them before 1980. Coakley et al. (1996, 1999) showed that when both saving rate and investment rate are difference stationary, the current account solvency implies the co-integration relationship between these variables. Then they used time series data from 23 OECD countries and 44 developing countries to conduct co-integration tests between savings and investment, and found that the co-integration relationship is more significant and stable in developed countries than in developing countries, which implies that the solvency constraint is binding for the developed countries.

Kydland and Prescott (1990) argued that the rate of technological change is related to the arrangements and institutions that a society uses and, more important, to the arrangements and institutions that people expect will be used in the future. Even in a relatively stable society like the U.S., there have been changes in the productivity growth in labor and capital since the Second World War. Since the underlying rate of productivity is hardly constant in Taiwan, the specification of linear time trend as the long-run trend is not appropriate. Here I propose an alternative trend removing regression equation:

$$z_t = a + bt + \mathbf{x}_t'\boldsymbol{\eta} + u_t, \tag{1}$$

in which z_t is either national saving rate or domestic investment rate at time t, t is the linear time trend, u_t is the cyclical component of z_t that deviates from its trend component $(a + bt + \mathbf{x}_t'\boldsymbol{\eta})$, $\boldsymbol{\eta}$ measures the effects of $\mathbf{x}_t$ on z_t, and $a, b, \boldsymbol{\eta}$ are parameters to be estimated. Here $\mathbf{x}_t$ is the $m \times 1$ vector of independent variables, contains the democracy index, financial openness, trade openness and other control variables, and m is the number of variables in $\mathbf{x}_t$.

According to (1), the short-run statistical relationship between savings and investment depends critically upon the variables selected in $\mathbf{x}_t$. One problem with this trend removing procedure is that difference between the original series and its trend component changes as more variables are included in $\mathbf{x}_t$.

Table 2: Trend-Removing Regressions

dependent variable / independent variable	constant	*t*	democracy	financial openness	trade openness	natural log value of populations	inflation rate
national saving rate	31.362***	0.113***	-0.154***	-0.201			
	(0.478)	(0.018)	(0.018)	(0.163)			
	27.334***	0.088***	-0.127***	-0.499***	0.054***		
	(1.319)	(0.022)	(0.022)	(0.165)	(0.018)		
	-564.863***	-0.184**	-0.066**	0.530**	0.102***	61.007***	-0.022
	(145.654)	(0.073)	(0.030)	(0.236)	(0.025)	(14.961)	(0.049)
domestic investment rate	30.654***	-0.151***	0.115***	0.129			
	(0.916)	(0.030)	(0.022)	(0.292)			
	12.509***	-0.266***	0.237***	-1.211***	0.242***		
	(2.566)	(0.029)	(0.025)	(0.264)	(0.035)		
	-601.881***	-0.470***	0.253***	-0.035	0.223***	63.472***	0.293***
	(145.292)	(0.065)	(0.023)	(0.261)	(0.028)	(14.901)	(0.043)

Note：1. National saving and domestic investment rates are in units of percentage.

2. Number in parenthesis is standard deviation of parameter estimate.

Apparently, there is a trade-off between the number of variables included and the time-series properties of u_t. Since the focus here is how democracy and financial openness affect the trend properties of national saving and domestic investment rates, the benchmark specification for (1) only includes the democracy index and financial openness in $\mathbf{x}_t$.

The results of the benchmark regressions in Table 2 indicate that democratic transition has significantly negative effects on the long-run trend of the national saving rate (coefficient estimate is -0.154), while it has significantly positive effects on that of the domestic investment rate (coefficient estimate is 0.115). The coefficient estimates for t in both benchmark regressions are positive and statistically significant. However, when I add the natural log of total population as an additional control variable in regression equation for savings, I find that the sign of the linear time trend coefficient estimate becomes negative.

During thte democratic transition period, the Taiwanese government initiated the national health care system and set up the national pension program as ways of responding to the needs of the general public. Apparently, social insurances lower the need of savings of the private agent and households for the precautionary purpose. This gives the negative relationship between democratic transitions and the long-run trend of savings. The channel influence of democracies and investment described in the introduction yields the positive relationship between democracy and the long-run trend of domestic investment rate.

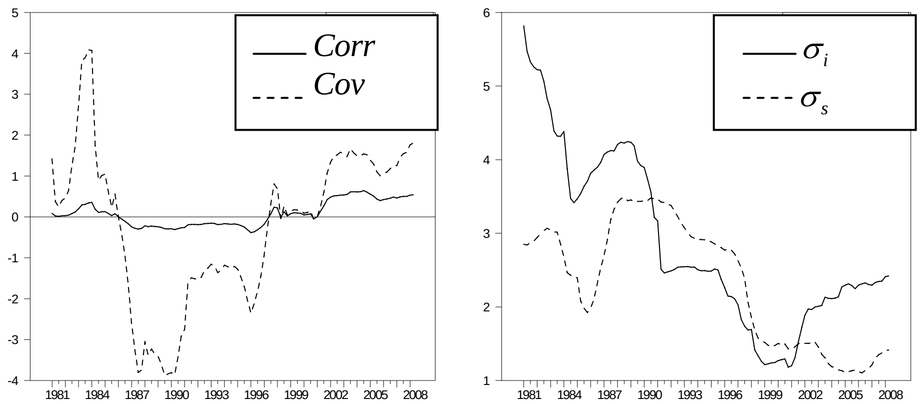
Financial openness has differential impacts on the saving and investment rates, but the coefficient estimates are not statistically significant. Financial openness can promote diversification of the industrial structure and the specialization of international production according to the country's comparative advantages and hence raise domestic investment. On the other hand, Feldstein (1994) argued that the firms' outbound FDI replaces their domestic investment almost by an equal amount. As displayed in Figure in Appendix, the measure of financial openness is highly correlated with the outbound FDI in Taiwan since 1992. Then a higher degree of financial openness due to an increase in outbound FDI leads to a decrease in domestic investment. Therefore, the net effect of financial openness on the long-run trend of investment rate depends on which channel of influence dominates. The regression result here indicates that neither channel dominates.

Fischer (2003) has stressed that financial openness has a positive effect on the long-run economic performance only when the country has a stable macroeconomic environment (good fiscal discipline, independent central bank, less capital controls). Coefficient estimates for trade openness are statistically significant and positive in the savings and investment trend removing equations. These results are consistent with the finding in Kose et. al (2003b). Finally, inflation rate has an insignificant effect on the saving rate, but has a significantly positive effect on investment rate. This suggests that stable macroeconomic environment help the formation of domestic capital. To satisfy the stationary requirement, the residuals obtained in the benchmark regressions will be used as the data series of the national saving rate, denoted s_t, and the domestic investment rate, denoted i_t.

In Section 4, I will use the DCC GARCH-X model to investigate how democratic transitions affect the correlation of s_t and i_t. Here I show figures with rolling statistics in Figure 2. To generate the numbers in Figure 2, I re-compute standard deviations of national saving and domestic investment rates and the covariance and correlation coefficient between these two variables for the fixed 40-quarter sample length which is shifted through the whole sample period. The rolling statistics are relevant for the question of how variable the statistics are across different sub-samples. When reading these figures, note that it is the end of the window that is displayed on the horizontal axis. For example, the first observation on the horizontal axis, which is dated 1980:4, is thus the statistic for the sub-period 1971:1 to1980:4.

The rolling standard deviations presented for the national saving and domestic investment rates in Figure 2 exhibit very different patterns of movement. The rolling standard deviation for savings (σ_s) slightly declined before 1985, then started to climb and reached the peak in 1988. After that, it declined until 2000. According to Figure 1, we also observed volatile variation in the balance on current account as shares of GDP during the period 1985 to 2000. On the other hand, the rolling standard deviation for domestic investment rate (σ_i) exhibited a very clear pattern of decline in the period from 1971 until 2001. After 2001, it returns to a mild inclining pattern. Except for the period 1990:4 to 2001:2, the values of σ_i are greater than those of σ_s, which is consistent with the finding in aggregate fluctuations that investment moves with grater volatility than savings.

Figure 2: The 10-year rolling standard deviations, covariance and correlation



For the rolling covariance and correlation coefficient between s_t and i_t in Figure 2, there are wide gaps between them before 1997. This indicates that the movements of both σ_i and σ_s are the dominant factors in the determination of the movement of the correlation between s_t and i_t. If democracies lower standard deviations of savings and investment, we should expect to see that the democratic transition lowers the correlation through decreasing variances of s_t and i_t. After 1997, variations between covariance and correlation coefficient shrunk and standard deviations of s_t and i_t returned to a stable pattern of variation and stayed at lower values, the importance of the movement of standard deviations in the determination of the movement of correlation decreased. The link between democracy and correlation through the channel of standard deviations is therefore weaker. Tesar (1991, Table 4) used data in 24 OECD countries to compute 5-year average of standard deviations for saving rate, investment rate and the balance on current account as shares of GDP. The statistics suggest that, for large-size countries, the lower the standard deviations are, the higher is the correlation. However, the statistical evidence for Taiwan is not due to the size of countries.

3. A DCC GARCH-X Model

To investigate how democratic transitions affect the variances of s_t and i_t and the covariance between these two series, I modify Engle (2002)'s DCC GARCH model to include the democracy index and other control variables as predetermined variables in the system. The model has time varying conditional variances, covariances and correlations. One major difficulty in modeling the conditional second moment for savings and investment is to keep the estimated variance and conditional covariance matrix positive definite in each period, when the democracy index and other control variables are included.

Assume that $c_t|\Omega_{t-1} \sim N(O, H_t)$, in which $\mathbf{c}_t = [s_t, i_t]$ is the 2×1 residual vector obtained in (1), Ω_{t-1} is the information set containing all the information available up to the time $t-1$, $\mathbf{H}_t$ is the conditional covariance matrix. Let h_{jt} denote the conditional variance of variable j at time t, for $j = s, i$. To assure the positive h_{jt}, I use an exponential GARCH-X model proposed in Lin and Yeh (2009) to capture the effect of $\mathbf{x}_t$ on the time varying conditional variances of s_t and i_t:

$$\log h_{jt} = \alpha_j + \beta_j \left| c_{j,t-1} / h_{j,t-1}^{0.5} \right| + \gamma_j \log h_{j,t-1} + \mathbf{x}'_{t-1}\boldsymbol{\pi}_j, \quad j = s, i, \tag{2}$$

in which α_j, β_j and γ_j are parameters to be estimated, and the $m\times1$ vector $\boldsymbol{\pi}_j$ measures the effects of x_{t-1} on $\log h_{jt}$ at time t.

Next, let $\mathbf{D}_t$ be the 2×2 diagonal matrix at time t, in which diagonal elements are $\sqrt{h_{st}}$ and $\sqrt{h_{it}}$. Then $\mathbf{c}_t$ can be standardized as $\mathbf{v}_t = \mathbf{D}_t^{-1}\mathbf{c}_t$. The covariance matrix of $\mathbf{v}'_t = [v_{st}, v_{it}]$ is given by

$$\mathbf{R}_t = \begin{bmatrix} 1 & \rho_t \\ \rho_t & 1 \end{bmatrix},$$

in which ρ_t is the conditional covariance between v_{st} and v_{it}.[1] $\mathbf{R}_t$ is also a correlation matrix for v_t. $\mathbf{H}_t$ and $\mathbf{R}_t$ satisfy $\mathbf{H}_t \equiv \mathbf{D}_t\mathbf{R}_t\mathbf{D}_t$.

As in Engle (2002), $\mathbf{R}_t$ can be obtained as follows. First, assume that the 2×2 conditional covariance matrix $\mathbf{Q}_t$ has the following structure:

$$\mathbf{Q}_t = \mathbf{W}(1-A-B) + A\cdot\mathbf{v}_{t-1}\mathbf{v}'_{t-1} + B\cdot\mathbf{Q}_{t-1} + \mathbf{C}\cdot\mathbf{x}'_{t-1}\boldsymbol{\psi}, \tag{3}$$

in which $\mathbf{W}$ is the sample covariance matrix of $\mathbf{v}_t$, $\mathbf{C}$ is the 2×2 matrix with 0 as diagonal elements and 1 as non-diagonal elements, A and B are parameters to be estimated, and the $m\times1$ vector of coefficients $\boldsymbol{\psi}$ measures the effects of $\mathbf{x}_{t-1}$ on $\mathbf{Q}_t$ at time t. If $\mathbf{C}\cdot\mathbf{x}'_{t-1}\boldsymbol{\psi} = \mathbf{0}$ and coefficients $(1-A-B)$, A and B are all positive, then Lemma 1 in Ding and Engle (2001) yields positive definite of $\mathbf{Q}_t$. However, if $\mathbf{C}\cdot\mathbf{x}'_{t-1}\boldsymbol{\psi} \neq \mathbf{0}$, equation (3) will not guarantee the positive definiteness of $\mathbf{Q}_t$ without additional restrictions. Letting $\mathbf{Q}_t^* \equiv \mathbf{W}(1-A-B) + A\cdot\mathbf{v}_{t-1}\mathbf{v}'_{t-1} + B\cdot\mathbf{Q}_{t-1}$, equation (3) becomes $\mathbf{Q}_t = \mathbf{Q}_t^*\left[\mathbf{I} + \mathbf{Q}_t^{*-1}\left(\mathbf{C}\cdot\mathbf{x}'_{t-1}\boldsymbol{\psi}\right)\right]$. According to Roger and Charles (1985, p. 468), if $\mathbf{Q}_t^*$ is positive definite and all the eigenvalues of $\mathbf{Q}_t^{*-1}\left(\mathbf{C}\cdot\mathbf{x}'_{t-1}\boldsymbol{\psi}\right)$ are greater than -1, then $\mathbf{Q}_t$ will be positive definite. Once $\mathbf{Q}_t$ is obtained, we have $\mathbf{R}_t = diag\{\mathbf{Q}_t\}^{-1/2}\mathbf{Q}_t diag\{\mathbf{Q}_t\}^{-1/2}$. Here the diagonal elements of $diag\{\mathbf{Q}_t\}$ are the corresponding diagonal elements in $\mathbf{Q}_t$.

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[1] In Bollerslev (1990)'s CCC GARCH model, ρ_t is assumed to be a constant.

Given the sample size of T, the biariate DCC GARCH-X model can be formulated as the following log likelihood function, apart from some initial conditions:

$$L_T = -\frac{1}{2}\sum_{t=1}^{T}\left(2\log(2\pi)+\log\left|\mathbf{H}_t\right|+\mathbf{c}_t'\mathbf{H}_t^{-1}\mathbf{c}_t\right)$$
$$= -\frac{1}{2}\sum_{t=1}^{T}\left(2\log(2\pi)+2\log\left|\mathbf{D}_t\right|+\log\left|\mathbf{R}_t\right|+\mathbf{v}_t'\mathbf{R}_t^{-1}\mathbf{v}_t\right),$$

which can simply be maximized over the parameters of the model. Without the assumption of conditional normality of $\mathbf{c}_t$, the estimator will still have the Quasi-Maximum Likelihood interpretation. Letting $\mathbf{h}_t' = [h_{st}, h_{it}]$, the log likelihood function can be written as the sum of the volatility part and the correlation part. The log likelihood function for the volatility part is

$$L_{hT} = -\frac{1}{2}\sum_{t=1}^{T}\left(2\log(2\pi)+2\log\left|\mathbf{D}_t\right|+\mathbf{c}_t'\mathbf{D}_t^{-2}\mathbf{c}_t\right),$$

which can be further decomposed into the sum of the two log likelihood functions for h_{jt}, $j = s, t$:

$$L_{hT} = -\frac{1}{2}\sum_{t=1}^{T}\sum_{j=s,i}\left(\log(2\pi)+\log(h_{jt})+c_{jt}^2/h_{jt}\right), \quad (4)$$

in which the volatility parameters to be estimated in (4) are simply those in (2): $\theta_j \equiv \left(\alpha_j, \beta_j, \gamma_j, \boldsymbol{\pi}_j'\right)$, $\quad j = s, i$. Equation (4) can be used to estimate the volatility parameters by separately maximizing each term. Since $\mathbf{c}_t'\mathbf{D}_t^{-2}\mathbf{c}_t = \mathbf{v}_t'\mathbf{v}_t$, it is straight from (4) that the log likelihood function for the correlation part is:

$$L_{RT} = -\frac{1}{2}\sum_{t=1}^{T}\left(\log\left|\mathbf{R}_t\right|+\mathbf{v}_t'\mathbf{R}_t^{-1}\mathbf{v}_t-\mathbf{v}_t'\mathbf{v}_t\right). \quad (5)$$

Here the correlation parameters to be estimated in (5) are those in (3): $\phi \equiv \left(A, B, \boldsymbol{\psi}'\right)$. The correlation part of the log likelihood function can be used to estimate the correlation parameters. The sum of L_{hT} and L_{RT} is L_T. Since the maximization of (4) over the volatility parameters does not depend on the correlation parameters in (5), we will adopt the two-step procedure to maximizing the log likelihood function. In estimation, we will use the sample estimate of covariance matrix obtained from (2) as the initial value for both $\mathbf{v}_t\mathbf{v}_t'$ and $\mathbf{Q}_t : \mathbf{W} = \mathbf{v}_0\mathbf{v}_0' = \mathbf{Q}_0$.

4. Empirical Results

Before estimation, we need to know whether or not σ_s, σ_i and Cov are time varying. Based on the Ljung-Box(1978) Q tests against the high-order serial correlation of s_t and i_t and the first-order serial correlation of the correlation between these two variables, Q(16) in Table 3 revealed that there are significant auto correlation in s_t, i_t and $s_t \cdot i_t$. Ljung-Box (1978) test statistics for autocorrelation in squared series ($Q^2(16)$) show that there are less autocorrelations in squared series. Then I use Engle (1982) ARCH test to detect if the squared residual series are serially correlated. The results is Table 3 confirmed that the squared series of s_t and i_t have serial correlation. Here, under the consideration of parsimonious representation, I choose EGARCH (p,q) model, and use traditional time series techniques applied to the squared series to select an EGARCH (1,1) structure for conditional variances. Finally, I use Tse (2000) test against the time-varying correlation between s_t and i_t:

$$\rho_t = \bar{\rho} + \delta \cdot s_{t-1}i_{t-1}.$$

When the null of $\delta = 0$ cannot be rejected, the Bollerslev (1990)'s GARCH model with constant conditional correlations should be used. Based on the Lagrange multiplier test statistic (16.65), the null is rejected at the 1% level of significance. Summing up, the tests discussed above do not present any serious evidence against the GARCH model with dynamically conditional correlation as a simple and parsimonious description of short-run dynamics for both national saving and domestic investment rates.

Table 3: Tests of GARCH Structure for s_t and i_t

Variable	skewness	kurtosis	JB test	Q(8)	Q(16)	$Q^2(8)$	$Q^2(16)$	ARCH(8)
s	0.477	1.332	16.779 (0.000)	243.650 (0.000)	316.332 (0.000)	166.902 (0.000)	184.767 (0.000)	74.480 (0.000)
i	0.709	3.544	91.069 (0.000)	145.503 (0.000)	181.690 (0.000)	62.971 (0.000)	79.306 (0.000)	34.940 (0.000)
$s \cdot i$	2.315	21.049	2903.215 (0.000)	40.136 (0.000)	49.576 (0.000)	–	–	–

Note: Number in parenthesis is the p-value of test statistics.

Table 3 also gives summary statistics of s_t, i_t and $s_t \cdot i_t$. It can be seen that all these three series all leptokurtic and the Jarque-Bera normalty test statistics show that these three variables are far from normality. Since s_t and i_t are serially correlated, I select the length of lag to be one to correct for autocorrelation in the covariance matrix of coefficient estimates.

4.1 Democracies and time-varying variances of savings and investment

Table 4 presents estimation results of the EGARCH(1,1) model for the volatility parameters. The coefficient estimates of α, β, γ in equation (2) for both s_t and i_t are all statistically significant under the restriction of $\boldsymbol{\pi}'_j = \mathbf{0}_{m\times 1}, \quad j = s, i$. The positive value of β states that the standardized residual has a positive effect on $\log h_{jt}$, while the positive value of γ signals the persistence in $\log h_{jt}$. These results are consistent with the test results given in Table 3. Allowing for the link between democracies (and financial openness) and the time-varying conditional variances of s_t and i_t, the values of the estimate of β and γ decrease. It suggests that democracies and financial openness can weaken the EGARCH(1,1) structures in conditional variances of s_t and i_t.

The coefficient estimates for the democracy index are positive in (2) for both s_t and i_t and statistically significant at the 1% and 5% levels, respectively. These results confirm that democracy can lower the conditional variances of both s_t and i_t. Democracy has a stronger negative influence on the volatility of s_t than on that of i_t. Recall that empirical evidence in Section 2 that the movement in the correlation between s_t and i_t before 1997 is mainly determined by changes in σ_s and σ_i. Apparently, the democratic transition affects the correlation through the following channel of influence. According to home biased effect, it is the less conditional variances induced by the democratic transition that motivate more private agents to put more savings in holding domestic assets relative to foreign assets.

These findings are much in line with previous results reported in the literature. For example, Lin and Yeh (2009) found that the falling volatility of real GDP, private consumption, and domestic capital formation in the 1980s can be explained by the democratic transition in Taiwan. Rodrik (1997) focused on real GDP, real consumption and investment. In each case, volatility is measured by calculating the standard deviation of annual growth rates of the relevant aggregate over the period 1960-1989. Then each measure of volatility is regressed on a number of independent variables, including the measure of democracy. He found that the coefficient estimated of democracy is negative and statistically significant in all cases. The association between the democratic transition and investment is strongest among the three aggregates. More recently, Rodrik and Wacziarg (2005) found, based on annual data from 24 countries with significant democratic transitions, that the volatility of growth rate of real GDP substantially declines after the democratic transition.

Table 4: Estimation of the volatility part

dependent variable / independent variable	σ_s^2			σ_i^s			σ_{CA}^2		
α	-0.108*	0.211	0.583	-0.304	0.198	0.452	-0.216*	0.321	0.204
	(0.061)	(0.324)	(0.621)	(0.208)	(0.273)	(0.473)	(0.127)	(0.286)	(0.541)
β	0.831***	0.710***	0.718***	0.579***	0.563***	0.565**	0.698***	0.574***	0.580***
	(0.116)	(0.195)	(0.180)	(0.089)	(0.110)	(0.125)	(0.093)	(0.107)	(0.100)
γ	0.599***	0.597***	0.590***	0.922***	0.727***	0.752***	0.847***	0.734***	0.724***
	(0.062)	(0.117)	(0.116)	(0.050)	(0.114)	(0.128)	(0.035)	(0.062)	(0.076)
democracy index		-0.861***	-0.894***		-1.078**	-0.979**		-0.779***	-0.800***
		(0.324)	(0.308)		(0.518)	(0.496)		(0.278)	(0.299)
financial openness		8.396	11.411		19.301*	20.129*		9.390	8.220
		(6.475)	(8.094)		(10.995)	(10.515)		(6.758)	(7.469)
trade openness			-0.435			-0.388			0.177
			(0.726)			(0.419)			(0.737)
likelihood ratio tests		14.561***	14.824***		25.638***	26.088***		15.869***	15.963***

Note:

1. Democracy index, financial openness and trade openness are all once-lagged variables.
2. Number in Parenthesis is the standard deviation of parameter estimate, and *,** and *** indicate 10%、5% and 1% level of significance respectively.
3. Critical values at 10%、5% and 1% level for $x^2(2)(x^2(3))$ statistic are 4.61（6.25）、5.99（7.81）and 9.21（11.34）.

Moreover, as Table 4 shows, the coefficient estimate of financial openness is positive and significant only at the 10% level in the EGARCH-X model for the conditional volatility of i_t, but the corresponding coefficient estimate is not significant even at the 10% level the conditional variance of s_t. Financial openness helps a country to diversify sources of funds in financing its domestic investment projects, but the volatility of domestic investment could increase as international funds respond more promptly to external shocks in countries with higher degree of capital mobility.

Thus, financial openness could spawn larger volatility in the investment rate. Prasad et al. (2004) divided 55 countries into two groups, according to the degree of financial openness in each country. Regression results indicate that countries in the group with higher degree of financial openness had larger volatilities of annual growth rates of real GDP, real domestic consumption, and real total consumption in the 1990s than those in the 1980s. Their finding is consistent with the results here.

Finally, the results of Table 4 suggest that democratic transitions also have negative effects on the balance on current account as shares of GDP. Note that the variance of the balance on current account as shares of GDP (σ_{ca}) is linked to the covariance between savings and investment through the following identity:

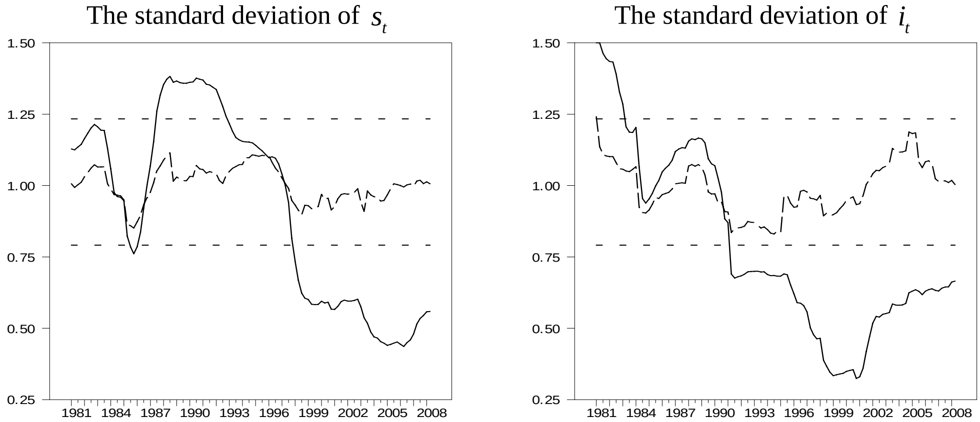
$$Cov = \frac{1}{2}\left[\sigma_s^2 + \sigma_i^2 - \sigma_{ca}^2\right].$$

Suppose that the democratic transition lowers variances of s_t and i_t, the identity implies a lower covariance between s_t and i_t (Cov), other things being equal. However, if the democratic transition also has a negative effect on the volatility of the balance on current account as shares of GDP, then the net effect of the democratic transition on the covariance becomes uncertain, depending on which effect dominating. On the other hand, given the estimation results in Table 4, financial openness has a positive and statistically significant effect on σ_i, but has negative and insignificant effects on σ_s. And we should expect to have a positive and significant effect of financial openness on Cov when we estimate the correlation part.

Figure 3 presents the rolling standard deviations of s_t and i_t with and without the democratic transition, financial openness, and GARCH effects being removed. To perform this analysis, I standardize the $\mathbf{c}_t$ series using the whole sample standard deviation and compute the 10-year rolling standard deviations for s_t and i_t. Then I use the $\mathbf{v}_t$ series to compute the 10-year rolling standard deviations. Variations in both rolling statistics show the democratic transition, financial openness, and GARCH effects on the volatility of s_t and i_t are evident. With these effects being removed, the rolling standard deviations generated by $\mathbf{v}_t$ stay within the 95% confidence band.

Given the above evidence, the bivariate DCC GARCH-X model seems to fit the data reasonably well. In order to assess the general validity of the model, a likelihood ratio test will be performed. Let L_{hT}^r be the natural log value of maximum likelihood function with the restriction $\boldsymbol{\pi}' = \mathbf{0}_{m\times 1}$ being imposed, and L_{hT}^u be the natural log value of maximum likelihood function without the restriction. The likelihood ratio test statistic is $\chi^2(m) = -2\left[L_{hT}^r - L_{hT}^u\right]$, in which the degree of freedom m is the number of predetermined variables in $\mathbf{x}_t$. According to the test result in Table 4, the null that the democratic transition and financial openness do not affect the conditional volatility of savings and investment is clearly rejected at the 1% level of significance. However, estimation result in Table 4 clearly reveals that the trade openness does not play such a significant role in accounting for variations in the volatilities of both s_t and i_t.

Figure 3: The 10-year rolling deviations



Note :

1. The parallel dotted lines indicate the 95% confidence band.
2. Solid line is the plot of rolling standard deviations generated by the series, while broken line is the plot of rolling standard deviations generated by the series.

4.2 Financial openness and the time-varying covariance between savings and investment

In this subsection, we use $\mathbf{v}_t$ obtained in the estimation of the volatility part as the data in the estimation of the correlation part to investigate the effects of the democratic transition, financial openness and trade openness upon the conditional covariance between s_t and i_t. Since the $\mathbf{v}_t$ series is serially correlated, I follow the procedure in subsection 4.1 and choose the lag length of autocorrelation to be one to correct the auto correlation in the covariance matrix of parameter estimates. Table 5 presents regression results.

Under the restriction of $\boldsymbol{\psi}' = \mathbf{0}_{m\times 1}$, estimates of A and B are all positive and statistically significant at the 1% level, which confirm that the conditional correlation between s_t and i_t has a DCC GARCH structure. Once the democratic transition and financial openness are taken into account in equation (3), the size of DCC GARCH effect decrease. It suggests that the democratic transition and financial openness can weaken the DCC GARCH structure in account for variation in the conditional correlation.

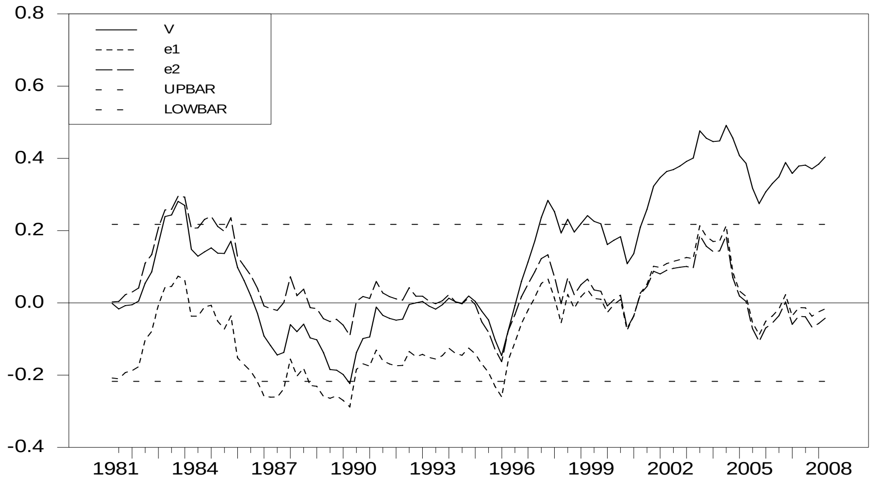
The coefficient estimate of the democratic transition in (3) is not statistically significant regardless of the specification of other predetermined variables. This is consistent with the implication in subsection 4.1. As discussed in the introduction, the effect of democratic transitions on the correlation between savings and investment depends on how democratic transitions affect the saving rate. Regression results in Table 5 suggest that the effect of the democratic transition on savings is uncertain and insignificant.

The coefficient estimate of financial openness is positive and statistically significant in (3), while the coefficient estimate of trade openness is negative and statistically significant. Evidently, an increase in the degree of financial openness inserts similar effect on both savings and investment. As displayed in figure in append, the measure of financial openness is closely associated with the outbound FDI in Taiwan since 1992. Does the outbound FDI rather than financial openness account for the variation in conditional covariance? To answer it, I replace the measure of financial openness with the capital outflow as shares of GDP and re-estimate EGARCH (1,1), and DCC GARCH-X model. The regression results using the two alternative measures of financial openness exhibit some quantitative difference, but display no qualitative difference. More specifically, higher ratio of the outbound FDI to GDP leads to a larger volatility of savings and investment, and a higher covariance. It appears that the channel of influence from financial openness on the covariance proposed by Feldstein did not receive strong supports from the data.

Table 5: Estimation of the correlation part

dependent variable / independent variable	*Cov*				
A	0.407*** (0.071)	0.388*** (0.039)	0.366*** (0.123)	0.360*** (0.076)	0.360*** (0.078)
B	0.287*** (0.103)	0.284*** (0.126)	0.286*** (0.095)	0.258*** (0.094)	0.259** (0.107)
democracy index		0.046 (0.039)	-0.171 (0.134)	-0.001 (0.125)	
financial openness			5.970*** (2.288)	7.937*** (2.282)	7.926*** (2.126)
trade openness				-0.205** (0.094)	-0.205** (0.080)
likelihood ratio tests		0.519	4.243	9.072**	9.072**

Figure 4: The 10-year rolling correlations using the v_t, e_{1t} and e_{2t} series

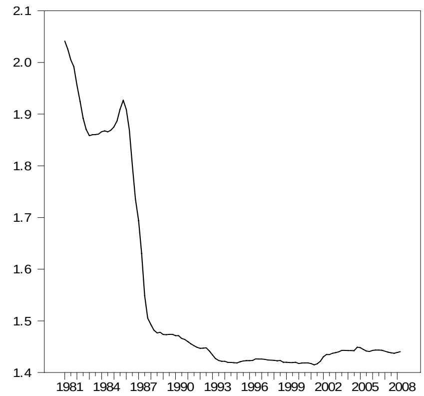


The coefficient estimate of trade openness is negative. Its statistical significance is weaker than the financial openness counterpart. Feldstein and Horioka (1980) argued that the correlation between savings and investment is weaker as the degree of trade openness is higher, since an open economy exposes itself to external shocks. Bahmani-Oskooee and Chakrabarti (2005) used annual saving-investment correlation data series from 126 countries over the period 1960-2000 and Sachs and Warner (1995)'s measure of economic openness to conduct panel regression, and reject the null that the saving-investment correlation does not depend on the degree of economic openness at 1% level of significance. Their finding is consistent with the finding here.

Next, I shall use the estimated conditional covariance ($\mathbf{H}_t$) to construct the orthonormal series from $\mathbf{c}_t$: $\mathbf{e}_t = \mathbf{H}_t^{-1/2}\mathbf{c}_t$. If the DCC GARCH-X model is properly specified, then the covariance matrix of $\mathbf{e}_t$ should be a 2×2 identity matrix. Further, according to the definition of $\mathbf{e}_t$ and $\mathbf{v}_t$, if $\mathbf{v}_t$ does not have a DCC GARCH-X structure ($A, B = 0$) and the democracy index, financial openness and trade openness do not have influence on the conditional covariance ($\boldsymbol{\psi}' = \mathbf{0}_{m\times 1}$), then $\mathbf{e}_t = \mathbf{v}_t$. Table 4 shows the rolling covariances and correlations constructed by $\mathbf{e}_t$ and $\mathbf{v}_t$. The dot lines in Table 4 are the 95% confidence band for the null of the zero correlation coefficient. Clearly, the rolling correlation generated by $\mathbf{v}_t$ rejects the null. Let $\mathbf{e}_{1t}$ denote the orthonormal series with the democracy index, financial openness and DCC GARCH effects being removed. Two findings are evident from the plot of $\mathbf{e}_{1t}$. First, the rolling correlation generated by $\mathbf{e}_{1t}$ is lower than that generated by $\mathbf{v}_t$, and has a weaker upward trend. Since the coefficient estimate of the democratic transition is not statistically significant in the correlation part, it is financial openness that accounts for a lower correlation before 2000. Second, the rolling correlations generated by $\mathbf{e}_{1t}$ are outside the 95% confidence band in the period 1986-1990. This is because that the democratic transition and financial openness cannot capture the volatile movement of the balance on current account as shares of GDP in this period. And it is this volatile movement that account for decreases in the saving and investment correlation.

Next, I compute the 10-year rolling correlations generated by the orthonormal series of random variable $\mathbf{e}_t$ with the trade openness being included as an additional variable in $\mathbf{x}_t$ denoted $\mathbf{e}_{2t}$. When comparing the two plots generated by $\mathbf{e}_{1t}$ and $\mathbf{e}_{2t}$, the plot generated by $\mathbf{e}_{1t}$ displays a narrower variation over the whole sample period. Both plots all stay within the 95% confidence band. Since trade openness has a negative effect on the conditional variance between savings and investment, the rolling correlation using $\mathbf{e}_{2t}$ are higher than those using $\mathbf{e}_{1t}$ before 1996 in particular. A narrower variation in the conditional correlation between $\mathbf{e}_{1t}$ and $\mathbf{e}_{2t}$ indicates that the trade openness has a weaker effect on the conditional correlation after 1996.

Given the regression results in the volatility part, the likelihood ratio test can be used to test against the null that predetermined variables in (3) fail to account for the time-varying conditional covariance. Let L_{RT}^{r} be the natural log value of maximum likelihood function with the constraint $\boldsymbol{\psi}' = \mathbf{0}_{m\times 1}$ being imposed, and L_{RT}^{u} be the natural log value of maximum likelihood function without the constraint. The likelihood ratio test statistic is $\chi^2(m) = -2\left[L_{RT}^{r} - L_{RT}^{u}\right]$. The test results in Table 5 confirm that financial openness and trade openness are significant determinants of variations in the conditional covariance, but the democratic transition is not.



To answer the question as to whether or not the democratic transition yields a higher correlation between savings and investment, I obtain another type of plot using the recursive technique of re-computing various statistics for an increasing length of sample period from the minimum length of 40 quarters to the full sample size, and inspect their convergence properties from the plot of the resulting series. Let β_{FH} be the theoretical value of Feldstein and Horioka (1980)'s saving retention ratio. Then β_{FH} and $Corr$ satisfy the following identity:

$$\beta_{FH} \equiv Corr \cdot \frac{\sigma_i}{\sigma_s},$$

All estimated β_{FH} and statistics in the above identity are obtained using recursive technique. The three findings from Figure 5 are as follows. First, the recursive estimates of β_{FH} reached the minimum value in 1987:2, stayed at negative values in the period between 1986 and 2001, and then exhibits an upward trend after 2000, which supports the notion that the democratic transition accounts for the difference of saving retention ratios between developing countries and developed countries. All the recursive estimates of β_{FH} are between -0.10 and 0.25, and not significantly different from zero. If the saving retention ratio indicates the degree of capital mobility, then this result suggests the short-run perfect capital mobility in Taiwan. Tesar (1991) used 10-year averaged national saving and domestic investment rates from OECD countries to estimate the saving retention ratio, and found that the saving retention ratio is significantly positive. Since OECD countries are democratic countries and have higher degrees of capital mobility than Taiwan, Tesar (1991)'s finding is consistent with evidence here. Second, β_{FH} is significantly higher than $Corr$ before 1987, which suggests that σ_i / σ_s is the dominant factor in the determination of changes in the saving retention ratio over the period 1971-1987. Further, β_{FH} and $Corr$ exhibit a highly coherent pattern of movement, and do not have wide variations between them after 1987. It implies that changes in $Corr$ not changes in σ_i / σ_s determine the changes in the saving retention ratio. Third and last, even though each individual standard deviation has a significant effect on $Corr$ through the channel of the democratic transition before 1997, the plot of σ_i / σ_s clearly indicates that the stable and low values of σ_i / σ_s explain why the saving retention ratio maintains a stable pattern. Since there was not a significant change in σ_i / σ_s after 2000, it is clear that a rising β_{FH} is due to a higher degree of financial openness in this period.

5. Concluding remarks

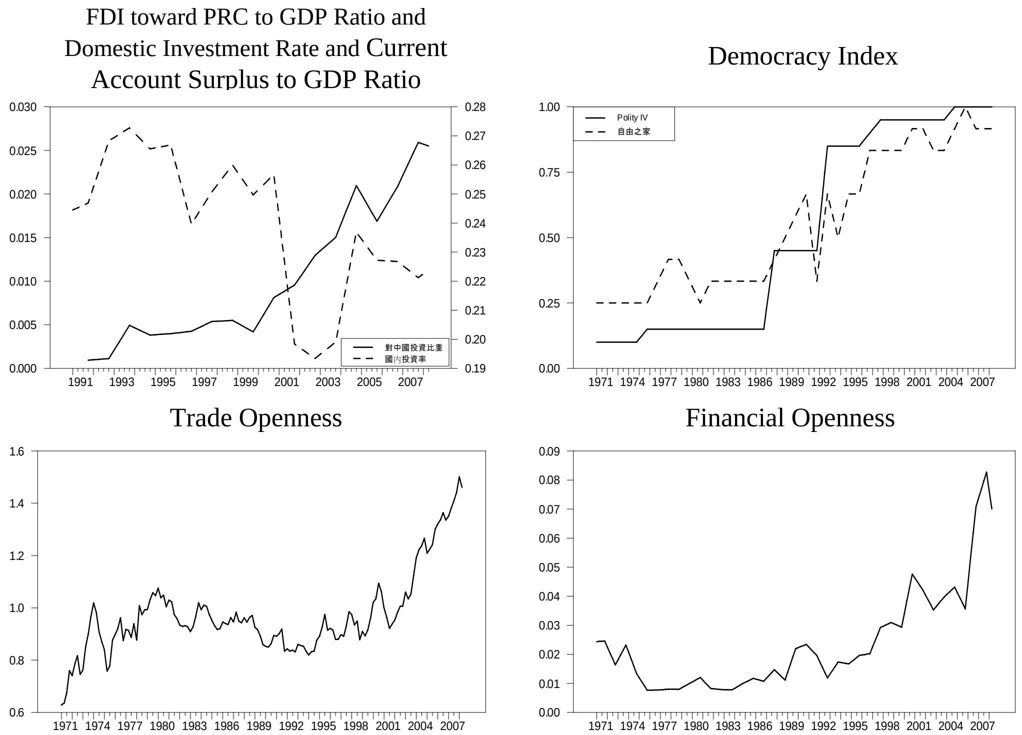
The purpose of this paper has been to investigate how democratic transitions affect the correlation between savings and investment in Taiwan. To achieve this goal, I first propose an alternative trend-removing procedure, based on the concept of long-run trends considered in Kydland and Prescott (1990), to induce the stationarity of national saving and domestic investment rates, and then introduce a DCC GARCH-X model to study the impact of democratic transitions on the correlation between savings and investment. I find that democratic transition lowers the correlation through decreasing variances of saving and investment, while financial openness raises it through increasing the covariance counterpart. Empirical findings suggest that democratic transitions are crucial to explaining why the correlation between savings and investment in developing countries is higher than that in developed countries. However, the question of how financial openness raises the covariance between savings and investment deserves further research.

Appendix: Data Definitions and Sources

Following Feldstein and Horioka (1980) and Obstfeld (1986), the amount of national savings is defined as gross national product (GNP) minus the sum of private consumption and government consumption (total consumption), the amount of domestic investment is the sum of gross domestic fixed capital formation and changes in stocks and work in progress, and the amount of balance on current account is defined as the level of net exports. These numbers divided by GDP give the national saving rate, domestic investment and the balance on current account as shares of GDP. The sample period is from the first quarter of 1971 (1971:1) to the second quarter of 2008 (2008:2). All the data are taken from "National Income Statistics in Taiwan".

Two measures of democracy are used in the paper: the democracy index published by Freedom House and Polity IV. I used the formula of (14-civil liberty-political right) /12 to convert the Freedom House index to the democracy index with the value between 0 and 1. I also use similar formula to convert Policy IV to obtain the similar index. According to Kose. et. al. (2003a), trade openness is defined as account of total trade as shares of GDP, and financial openness is defined as amount of capital inflows and outflows as shares of GDP.

Figure 5: Democracy indexes, financial openness and trade openness



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