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Trajectories to high income: comparing the growth dynamics in China, Korea, and Japan with cointegrated VAR models

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Trajectories to high income: comparing the growth dynamics in China, Korea, and Japan with cointegrated VAR models

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Trajectories to high income: comparing the growth dynamics in China, Korea, and Japan with cointegrated VAR models

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Abstract:

We analyze and compare the patterns of economic growth and development in China, Korea, and Japan in the post-war period. The geographical proximity and cultural affinity between the three countries, as well as the key role of the development state in the economies, suggest that an analytical comparison would be a meaningful and valuable exercise. Furthermore, Korea and Japan are two of the few economies that have jumped from middle income to high income in a short period and thus offer potentially valuable lessons for China. China is following a structural change that Korea and Japan underwent decades ago. We use Cobb–Douglas production functions to assess the long-run equilibrium relationships between per capita GDP, capital, and labor as well as the features of structural change by means of cointegrated vector autoregressive (CVAR) models. We show that such equilibrium relationships cannot be rejected for all three countries, while the evidence is stronger for China and Korea than for Japan. Our hypothesis tests show that the estimated Cobb–Douglas production functions display coefficients of capital and employment that sum up to one and broken linear trends that can be attributed to structural breaks and (changes in) total factor productivity (TFP) growth. We observe a striking similarity between the Korean and the Chinese experience, which gives some optimism that China may be capable of graduating to high income, like Korea.

Keywords: aggregate production function, comparative economic growth, China, Korea, Japan, economic development

JEL codes: E23, O47, O53, O57, P52

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1 Introduction

Since the introduction of market reforms in 1978, decades of world-topping economic growth have transformed China into the world's second-biggest economy and an upper-middle-income economy. China's remarkable economic transformation, triggered by a systemic shift from a centrally planned economy to a more market-oriented economy, may indeed be the most significant development in the global economic landscape since the Second World War. However, since the global financial crisis of 2008–2009, China's growth has slowed down visibly, although it continues to grow at a healthy pace. While the slowdown is partly due to a less benign external environment, it is largely due to structural factors, such as rebalancing towards domestic demand and consumption, rapid income convergence towards high-income countries, population aging, and tertiarization. China is already an upper-middle-income country with an income level at which growth typically slows down (see, for example, Eichengreen et al. 2012, 2014). Therefore, to some extent, the slowdown is a necessary transition to a more balanced and sustainable growth paradigm, not least against the many imbalances that built up during the high-growth decades.[1]

At the same time, there is no guarantee that China's transition from middle income to high income will be as smooth and fast as its transition from low income to middle income. In fact, economic theory suggests that sustaining rapid growth will be difficult because marginal returns to capital eventually decline as an economy grows richer and acquires a larger stock of capital. The gains from shifting workers from low-productivity agriculture to higher-productivity manufacturing also eventually decline. Furthermore, as countries approach the global technology frontier, they must begin to develop new technology on their own instead of relying exclusively on importing advanced technology from abroad. Generally, the essence of economic growth shifts from input accumulation – that is, deploying more capital, labor, and other inputs – to total factor productivity growth – that is, using all those inputs more efficiently.

Empirically, a large number of middle-income countries have failed to graduate to high-income status in a reasonable period of time. This well-known stylized fact has given rise to the concept of the middle-income trap. Of 101 middle-income countries in 1960, only 13 proceeded to high-income status by 2008.[2] Will China be able to follow in their footsteps? Of the 13 countries mentioned, only a few appear to be comparable with China. Albert et al. (2015) point out that only Japan, Korea, Taiwan, and Israel followed a growth strategy similar to that of China: export-led growth paired with strong investment. Of special interest and relevance to China is the experience of Japan and Korea, which are relatively large countries. Although both countries are nowhere near as large as China, they are much larger than the small city-states of Hong Kong and Singapore and substantially larger than Taiwan.

The central objective of our paper is to assess empirically China's prospects for transcending the middle-income range by taking a look in the rearview mirror and comparing China's past experiences with those of Japan and Korea. Accordingly, we analyze and compare the economic

[1] See Wagner (2017) and Maliszewski and Zhang (2015).
[2] See World Bank (2013).

growth and structural transformation trajectory of China, Korea, and Japan on the macroeconomic level by estimating Cobb–Douglas production functions in a multivariate cointegration framework. To identify our models, we use ex-ante knowledge about historical events as well as the results of Bai and Perron's (2003) structural break tests of the time series.
The rest of this paper is organized as follows. Section 2 reviews the relevant literature and describes the economic growth experiences of China, Korea, and Japan. The section divides the experiences of these countries into different stages of economic development and structural transformation. Section 3 derives our theoretical hypotheses for the equilibrium relationships and analyzes and compares the Chinese, Korean and Japanese growth experiences more rigorously based on econometric analysis performing structural break tests and cointegration analysis. Section 4 concludes the paper.

2 Literature

Will China be able to close the gap to the high income economies? Some convergence theories predict that this will be the case. Lee (2016) argues that China's growth experience overall resembles the experience of Japan and Korea. Figure 1 compares the status quo of China, Korea and Japan in terms of catching-up with the US.[3] Every value expresses GDP per Capita in relation to the US. Hence, the difference between the countries' graphs and the x-axis can be interpreted as the remaining catching-up potential.

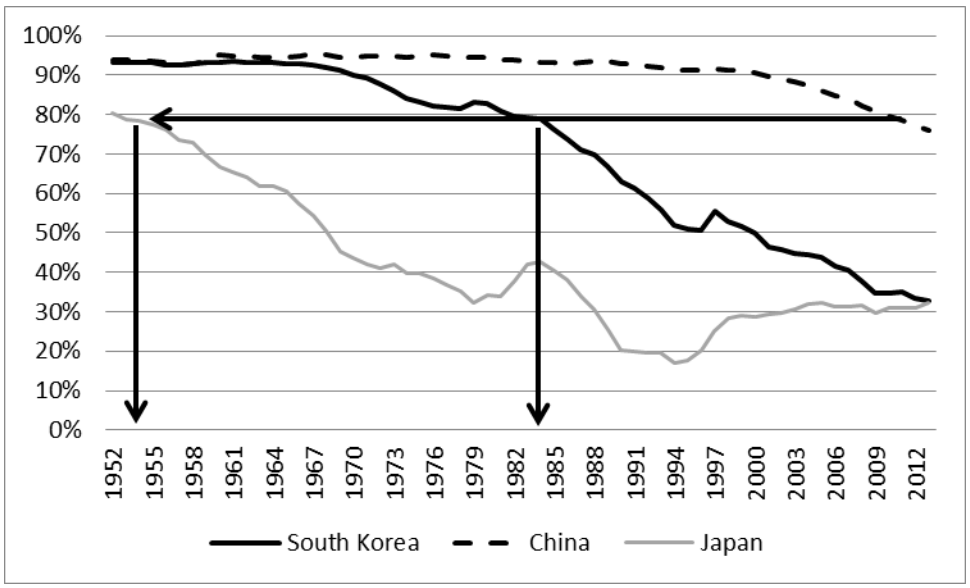

Figure 1 Catching-up potential of Japan, Korea and China to the US
Source: Own calculations with Penn World Table PWT version 9.0 data.

We can observe that all three countries are converging towards the US's GDP per capital level, yet the remaining catching-up potential is different for the 3 economies. Japan caught up quickly

[3] See Lee (2016) for a similar comparison, albeit with a different representation.

with the US and graduated to high-income status by the late 1960s.[4] After the nation-building phase in Korea had ended in the early 1960s, it grew dynamically until the early 1980s and then entered several stabilization processes, which are continuing until today. High-income status was reached around the turn of the millennium. By about 2014, Korea had caught up with Japan. Due to its specific circumstances, the fast-growth phase in China only started in the late 1970s. The remaining catching-up potential for China is still high. Based on Fig. 1, we can argue that the developments in each country lead those of the 'following' countries by about 27 to 30 years. The clear interpretation of this figure is that the 3 countries started from different initial conditions and thus have also undergone different phases of growth. Roughly, we can characterize these phases as follows.

China's post-war economic development can be divided broadly into three different stages: the centrally planned economy (1949–1977), the transition to a more market-oriented economy and economic take-off (1978–2000), and post-WTO globalization, technological upgrading, and transition towards a high-income economy (2001–present) (see Chow 2015).

Korea's post-war economic development can be divided roughly into four stages: nation building and the beginning of a modern market economy (1948–1959), economic take-off (1960s–1979), stabilization and transition toward a high-income economy (1980–1997), and restructuring, reinvention, and maturation (1998–present) (see Moon 2016).

Japan's post-war economic development can be divided into five different stages: a post-war reconstruction phase (1945–1954), a fast-growth phase (1955–1972), the oil shock phase (1973–1984), the bubble economy phase (1985–1990), the lost decade (1990–2000), and a revitalization phase (2004–present) (see Ifo Institute for Economic Research and Sakura Institute of Research 1997; Inkster 2001; Syed et al. 2009; Yoshino and Taghizadeh-Hesary 2015).

Looking at the drivers of the convergence observed in Fig. 1, the related literature points towards growth theories and structural change. Lewis (1954) argues that structural change in the sense of reallocation of labor between sectors is an important driver of economic growth. In Japan, the service sector share in employment surpassed the agricultural share in 1956 (the industry sector surpassed the agricultural sector in 1962); the corresponding intersections happened in 1980 (1984) in Korea and in 2011 (2014) in China. Thus, structural change in Japan leads that in Korea by 24 years and Korea's structural change leads China's by 31 years (Murach and Wagner 2017). Thus, the difference in the catching-up process seems to be mimicked by the respective process of structural change.

Dalgaard and Strulik (2013) merge unified growth theory and neoclassical growth theory to explain income differences between countries. They argue that the beginning of the fertility decline is an important indicator for the take-off of an economy. This take-off can be defined as the transition from a phase of economic stagnation to a phase in which these countries grow persistently in terms of income per capita. To time the take-off, two measures are proposed, the

[4] High income can be defined as > 50% of the US GDP (see Glawe and Wagner 2016).

first of which is the year of fertility decline. The point of fertility decline can be defined as the point at which the population growth decreases. The second measure proposed by Dalgaard and Strulik (2013) is the year of industrialization, which measures the first year when the industry share in the total employment is greater than that of agriculture. For Japan, Korea, and China, the initial years of fertility decline were 1950, 1960, and 1970 (as reported by Reher 2004). The year of industrialization corresponds to the years in brackets in the previous paragraph. Table 1 summarizes our findings.

Table 1 Summary of potential structural breakpoints

	Japan	Korea	China
Year of fertility decline (see Reher 2004)	1950	1960	1970
Structural change (see Murach and Wagner 2017)	1956 (1962)	1980 (1984)	2011 (2014)
Year of industrialization (see Dalgaard and Strulik 2013)	1962	1984	2014

Source: Own representation.

Dalgaard and Strulik (2013) argue that neoclassical growth theory is valid especially in the post-take-off phase. Neoclassical growth theory considers in particular the effects of input factors on economic growth. In this modelling framework, the Solow (1956) model explains growth as a function of capital, labor, and technological progress. The seminal work by Mankiw et al. (1992) empirically proves the general explanatory power of such models. They show that the standard Solow model indicates a share of 0.60 of physical capital in income, which, including a measure of human capital, decreases to 0.31.[5] Applying a different methodology, Barossi-Filho et al. (2005) find support for a capital share of one-third. By contrast, Hamilton and Monteagudo (1998) argue that physical capital may indeed have a greater impact on growth than in the original analysis by Mankiw et al. They assert that capital may also be able to include technological improvements. Besides the impact of physical capital, they find that population growth and the initial level of output are important in the analysis. Other works continue to extend this fundamental framework in different directions, promoting a differentiated view on the sources of economic growth, for example an open-economy model or the effects of health capital (see e.g. Barro et al. 1995; Knowles and Owen 1995). This augmented model type is also applied to China in recent research to evaluate the growth prospects of the economy (see Barro 2016; Lee 2016).

Besides cross-country comparisons like that undertaken by Mankiw et al. (1992), some authors focus solely on the growth dynamics in one or two countries. Durlauf et al. (2001) recommend considering local Solow models, as country-specific heterogeneity is likely to exist. This shows that examining the dynamics of individual countries can be a helpful exercise in analyzing the common drivers of economic growth. Chow (1993, 2015) analyzes the growth dynamics for

[5] See Mankiw et al. (1992, Tables I and II).

China. Chow and Lin (2002) compare the developments in Mainland China with those in Taiwan, arguing that the initial conditions were different and hence the analysis of the growth dynamics is an interesting exercise. Obviously, the initial conditions in which one country is nested are important and have implications for its future growth prospects.

Chow (1993) estimates production functions for the Chinese economy and its sectors. For this purpose, he constructs a measure of capital formation. He analyzes the effects of important political campaigns, like the Great Leap Forward (1958–1962) and the Cultural Revolution (1966–1976). Based on a figure that plots log (national income/labor) against log (capital/labor) for the period 1952–1985, he argues in favor of the exclusion of the years 1958–1969 because they are 'abnormal because of the great leap upheavals of the Great leap Forward movement and the Cultural Revolution'.[6] For the aggregate economy, capital coefficients of about 0.60 are estimated. A further finding of his paper is the absence of technological progress between 1952 and 1980. Chow explains this finding by pointing out that China concentrated on central planning and investments in heavy industries. Thus, incentives for private enterprises were scarce during this period. He argues further that technological progress is an important feature of market economies but is unlikely to be visible in an economy in which private initiatives and the adoption of new technologies from abroad are not present.

Cheremukhin et al. (2015) support this view, as they find in their review of the related literature that TFP growth accelerated moderately between 1978 and 1985. Of the industrial TFP growth in the 1980s, 87% can be explained by improved incentives, intensified product market competition, and improved factor allocation (see also Li 1997).

Chow and Lin (2002) perform a comparative analysis of Mainland China and Taiwan. For China, the sample is now 1952–1998. They estimate a Cobb–Douglas production function with a trend beginning with $t = 1$ in 1979. The coefficient of ln (K/L) and the time trend t are estimated with 0.64 and 0.0262 remaining unchanged over the sample period. In comparison with Taiwan, Chow and Lin comment on the relatively small exponent of labor in China. This is explained by the relative abundance of labor in Mainland China. Chow and Lin argue that, over time, labor will be less abundant as the economy grows and the ratio of capital to labor increases. They add that, in their sample, the abundance of labor has persisted, which is explained by still-poor regions in Western China with lower wages.[7]

In a third approach, Chow (2015) investigates the Cobb–Douglas production function for China between 1952 and 2012. For this sample, the coefficient of ln (K/L) is estimated as 0.59 and the trend is estimated as 0.298. Chow basically reproduces the arguments of Chow and Lin (2002) to support his results. He additionally argues that Mankiw et al. (1992) estimate a coefficient of 0.6

[6] Fig. 2 in Section 3.2, which plots log (income/population) against log (capital/labor) displays the same pattern of non-linearity of the years 1958–1969 (see also Chow 1993, Fig. I, p. 821).

[7] It is noteworthy that the comparison between Taiwan and Mainland China is based on different estimations. For Taiwan, human capital is considered, which reduces the coefficient of various capital stock measures in relation to labor to about one-third. This is consistent with the results obtained by Mankiw et al. (1992).

with data for many developing countries. The analyses by Chow will be our roadmap for our estimations in Section 3.

Finally, we would like to comment briefly on some similar inner dynamics and characteristics of the three countries that we investigate to support the similarity of the growth dynamics. Morck and Yeung (2017) review the financial and economic developments in East Asia, focusing on Japan, Korea, and China. As common characteristics of the three countries, they identify, among others, the following. i) A weakening in traditional institutions is observed before the phases of rapid economic development.[8] Thus, a serious obstacle to growth-generating reforms was reduced. ii) Literacy in the three countries was high early on, which was helpful for building up human capital later. iii) Moreover, relative openness due to the export- and investment-led growth model brought about early international competition and supported the state-directed process of foreign technology adaption. iv.) Product diversification was reached via business groups, like the zaibatsu in Japan, chaebols in Korea, and state control over huge sectors of the economy in China. v) Lastly, Japan's zaibatsu, Korea's chaebols, and China's state machinery were able to create a 'Big Push'.

This seems to be enough to motivate a close look at the production function of the three economies in comparison.

3 Empirical analyses

3.1 Developing theoretical hypotheses

As our analysis is motivated by the works of Chow (1993, 2015), we orientate ourselves towards his procedure. Chow (2015) suggests a Solow (1956)-type model of economic growth. In this framework, one equation explains the aggregate output through capital and labor inputs in the form of a Cobb–Douglas production function:

$$Y_t = A \cdot e^{\gamma t} \cdot {K_t}^{(1-\alpha)} \cdot {L_t}^{\alpha} \tag{1}$$

The exponents $(1-\alpha)$ and α measure the rate of change of output vis-à-vis changes in capital K_t and labor L_t. γ measures the change in output per unit over time even if the inputs of capital and labor are not changed. This may be the consequence of institutional or technological change due to economic reform policies. For $\gamma = 0$, this change is not present. If $\gamma = 0.01$, this can be interpreted as TFP growth of 1% per year. A can be interpreted as the initial level of technology in the economy (see Chow 2015).

As we are especially interested in the evolution of the *per capita* GDP, we have to include the following changes. We would like to have ${Y_t}/{N_t}$ on the left-hand side of equation (1). Thus, we

[8] This is not necessarily identical to the definition of the fast-growth phases earlier in this section.

divide (1) by N_t on both sides. However, on the right-hand side, we would rather like to express N_t by different variables to avoid a direct source of collinearity.

Hence we express N_t by L_t: $N_t = G_t \cdot L_t$. With $G_t \geq 1$. This is the inverse of the employment rate.

This gives

$$\frac{Y_t}{N_t} = \frac{A \cdot e^{\gamma t} \cdot K_t^{(1-\alpha)} \cdot L_t^{\alpha}}{G_t \cdot L_t} \tag{2}$$

or

$$G_t \cdot L_t \cdot Y_t = N_t \cdot A \cdot e^{\gamma t} \cdot K_t^{(1-\alpha)} \cdot L_t^{\alpha} \tag{3}$$

Taking natural logarithms we obtain

$$g_t + l_t + y_t = n_t + a + \gamma \cdot t + (1-\alpha)k_t \cdot + \alpha \cdot l_t \tag{4}$$

Rearranging terms gives

$$y_t - n_t = a + \gamma \cdot t + (1-\alpha)k_t \cdot + \alpha \cdot l_t - l_t - g_t \tag{5}$$

Collecting terms gives

$$y_t - n_t = a + (1-\alpha)k_t - (1-\alpha)\, l_t - g_t + \gamma \cdot t \tag{6}$$

This leaves us with a linear equation with testable hypothesis for the coefficients. Specifically we will estimate and test:

$$(y_t - n_t) = a + (1-\alpha) \cdot (k_t - l_t) + \gamma \cdot t - g_t + \varepsilon_t \tag{7}$$

3.2 Data Description & Historical Events

We use Penn World Tables (PWT) of the version 9.0 for our estimations. For China, Korea and Japan we collect data for gdp, capital, employment and population. Data are available from and 1952 until 2014 for China, from 1953 until 2014 for Korea, and from 1950 until 2014 for Japan. Chow (2015) constructs capital data for China from publications of the China Statistical Yearbook. However, Holz (2006) discusses an earlier version of Chow's data set and lists a number of shortcomings of Chow's measures. In a recent publication, he proposes some newly constructed capital measures for China (see Holz and Yue 2017). We do not try to construct capital measures for China, Korea, and Japan ourselves but instead rely on the time series provided by the PWT (see Feenestra et al. 2015). This naturally causes difficulties, as research shows that conclusions drawn under one version of the PWT may not hold under another version (see Ponomareva and Katayama 2010).

Fig. 2 displays prima facie evidence of the relationships between capital intensity and per capita GDP in China, Korea, and Japan. The triangles in dark grey refer to data from China, the diamonds in dark grey refer to Korea, and the squares in light grey represent data for Japan.

The relationships for China and Korea especially appear to be fairly linear and very similar. Japan seems to present about the same slope as Korea and China until about 1973. From a relative perspective, China seems, for an extended period, to follow a path that is very close to the previous Korean experience. Japan seems to be a different case, presenting much less linearity and a different average slope. At the beginning of China's development path (around the years 1958–1969), we observe the same irregularities as Chow (1993). That is why Chow (2015) excludes these data from his estimations.[9]

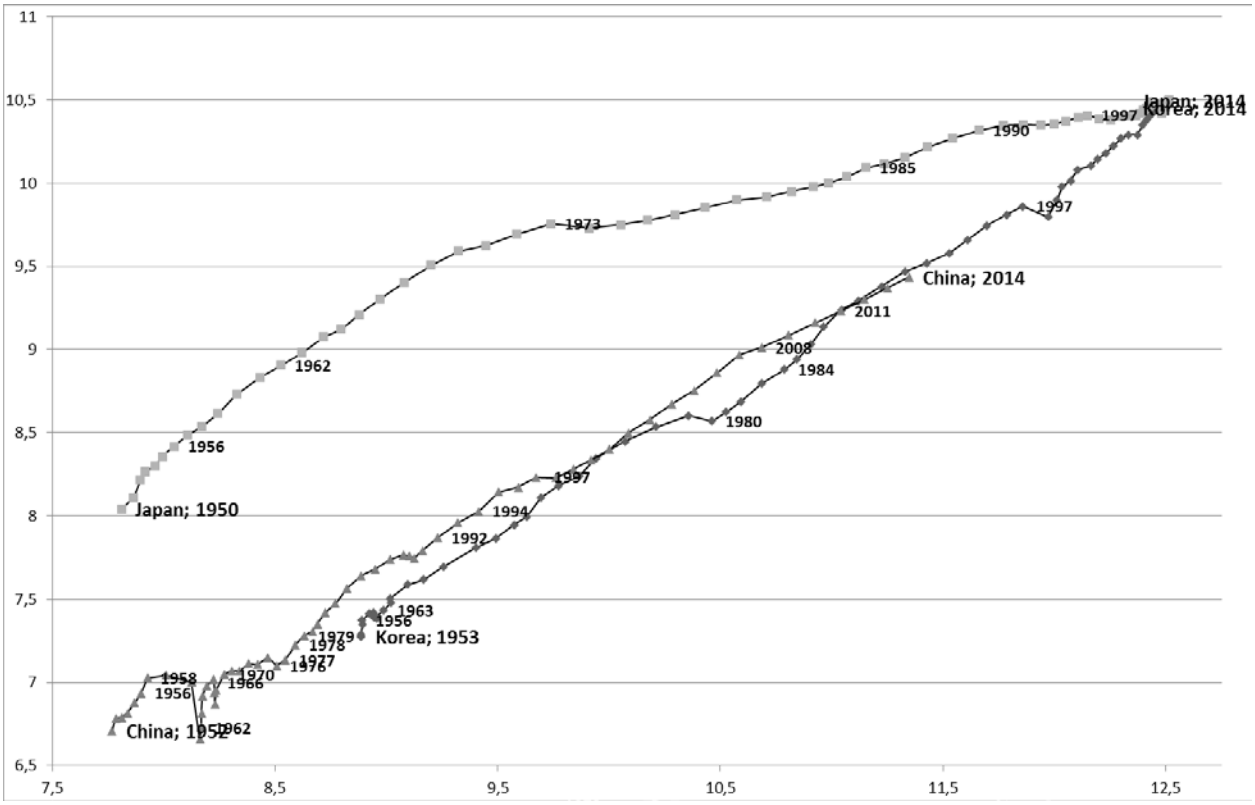

Fig. 2 Growth trajectories of Japan (top left), Korea (bottom left), and China (middle); the x-axis is the logarithms of capital intensity (K/L), and the y-axis is the logarithms of per capita GDP (Y/N)
Source: Own calculation with PWT 9.0 data.

A more detailed view of the individual country time series is presented in Figs 3 to 5. Fig. 3 shows that China's per capita growth was comparably volatile before 1979. The same is true for the growth in the capital–employment ratio, which became increasingly strong over the sample period, with the exception of a strong slump around the end of the 1980s.[10] The inverse of the

[9] See also Section 2.
[10] This is also in line with the assessment by Chow and Lin (2002) in Section 2.

employment rate decreased over the whole sample period, indicating that the overall employment in the population increased.

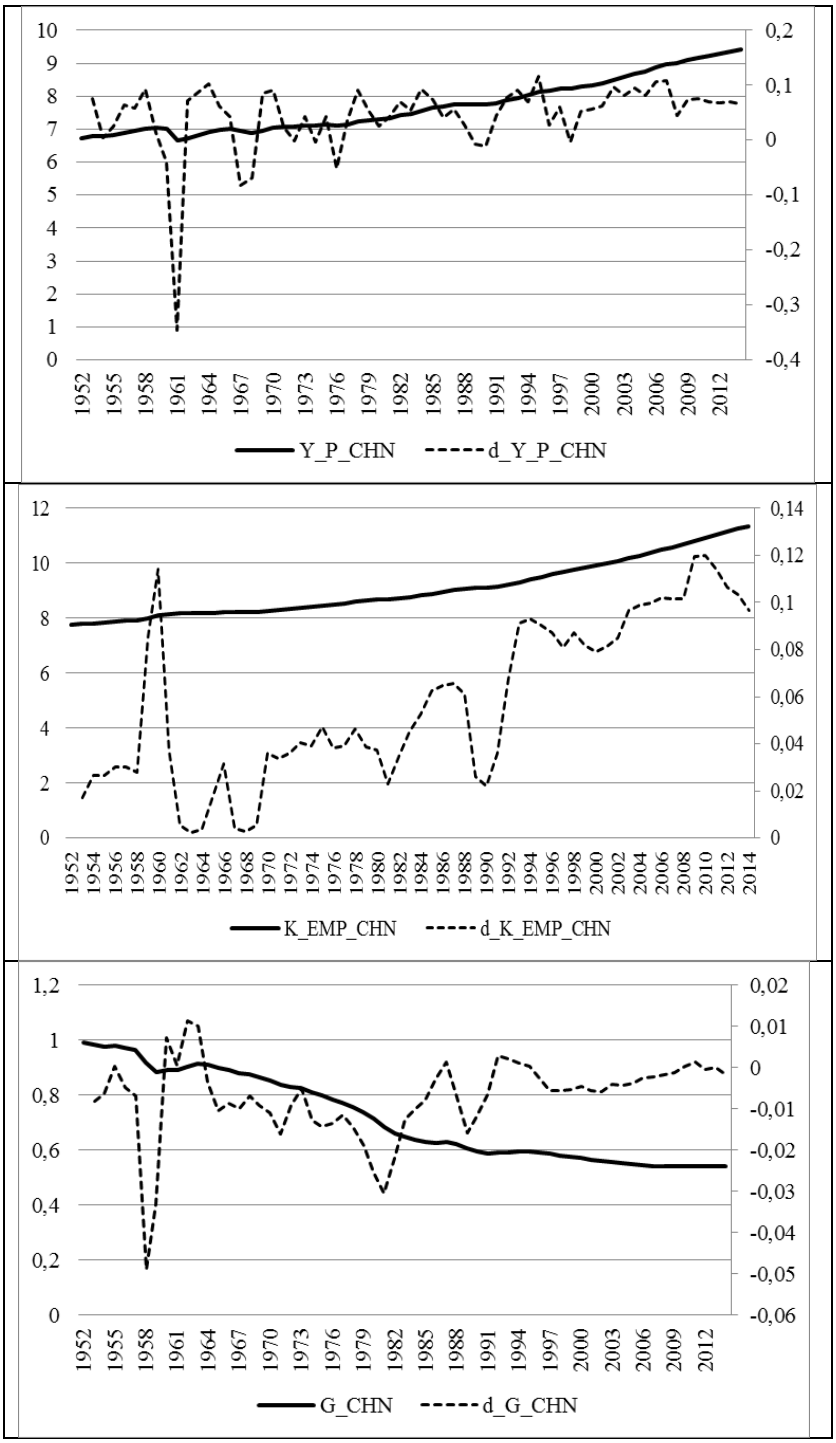

Fig. 3 China: GDP per capita, capital–employment ratio, and inverse of the employment rate
Source: PWT 9.0; own calculations. Logarithms of levels and first differences.

Fig. 4 displays the developments in Korea. The per capita GDP growth rates peaked around 1980 and then decreased. Important recessions happened in 1980 and during the Asian Crisis.

Altogether, the per capita GDP growth rates display a humped-shaped pattern. The capital–employment ratio rose strongly in the 1960s (the fast-growth phase), and the growth rates were very high until the 1980s, when they experienced a strong cutback and decreased further throughout the Asian Crisis. The inverse of the employment rate shows the same behavior as in China.

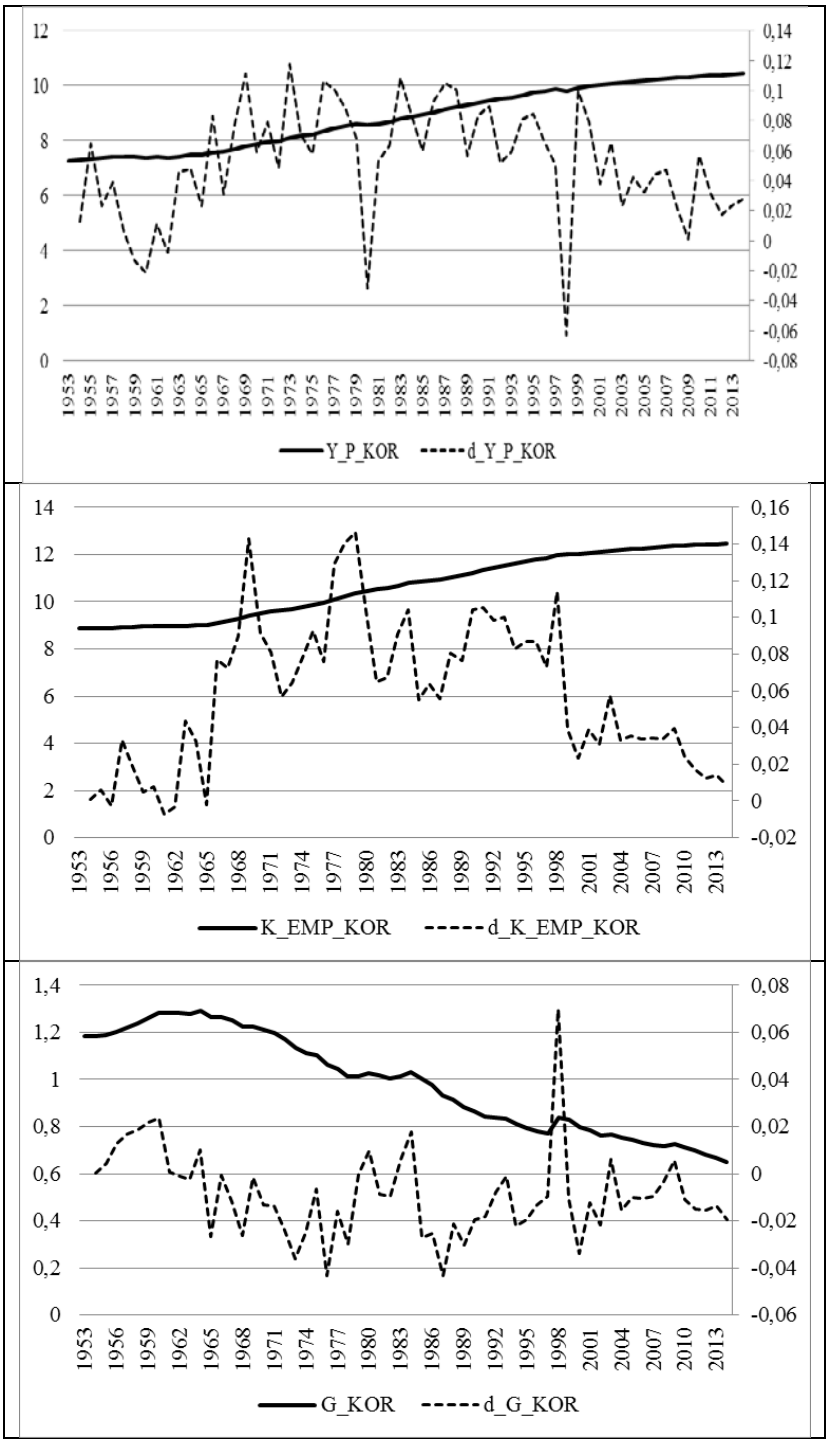

Fig. 4 Korea: GDP per capita, capital–employment ratio, and inverse of the employment rate
Source: PWT 9.0; own calculations. Logarithms of levels and first differences.

In Japan (see Fig. 5), we observe a more or less stable decrease in the per capita growth rates over the whole sample period. Higher per capita growth rates corresponded to fast increases in capital intensity until the oil crisis in 1973. Afterwards, the capital–employment ratio decreased, although we observe a short-lived boom during the second half of the 1980s. The inverse of the employment rate rose until the oil crisis and then seems to have followed the Japanese business cycle. Fig. A1 in the Appendix shows the separate time series for capital and employment for the three countries.

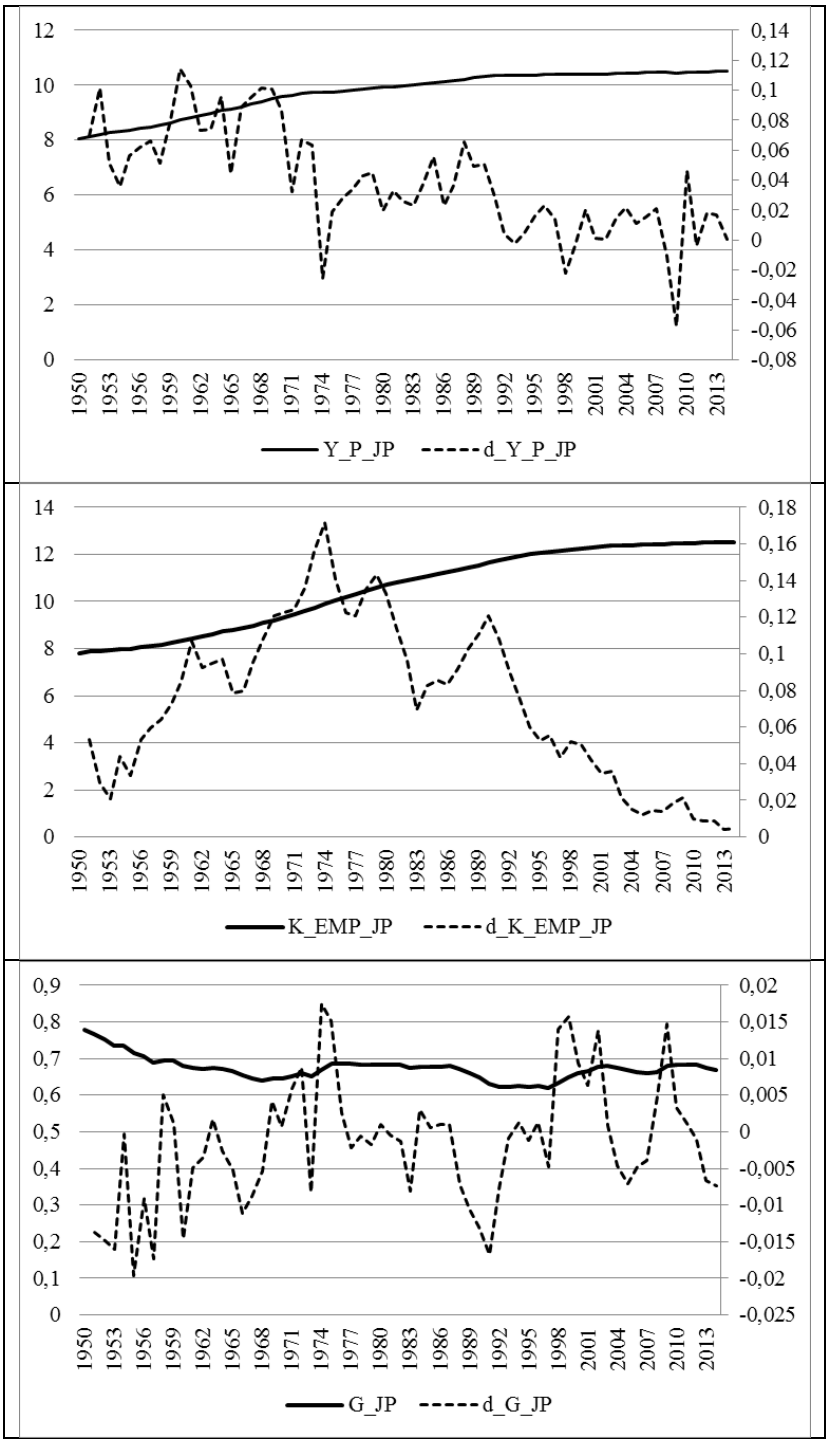

Fig. 5 Japan: GDP per capita, capital–employment ratio, and inverse of the employment rate
Source: PWT 9.0; own calculations. Logarithms of levels and first differences.

3.3 Univariate properties of the data

For the unit root tests, we analyse the period from 1956 until 2014. In the previous section, we saw that the growth rates decreased or increased over time. We thus assume that the time series

under consideration present deterministic trends and hence test especially for unit roots of the most general form, allowing for a constant and a deterministic trend. According to the unit root tests, the per capita GDP can be assumed to be integrated of order 1 (I(1)) for all three countries. The same is valid for the employment time series. The tests lead to less convincing results for the capital and capital–employment ratio series. We conclude that the levels of the time series are at least integrated of order one (I(1)). The capital and capital–employment ratio time series may even be I(2).[11]

The I(2) properties of the capital–employment ratio time series could require an I(2) analysis. However, Juselius (2006) points out that the behavior of an I(2) trend can be mimicked with an I(1) stochastic trend around a broken linear deterministic trend. Thus, it is possible to avoid an I(2) analysis by including an appropriate number of deterministic (linear, broken linear, or quadratic) trends in the data. This is the procedure that we choose for this section. To obtain a better understanding of where deterministic components could be fitted into our models, we analyze the time series for potential structural breaks in the next section.

3.4 Structural break tests

From a visual inspection of the time series growth rates, we can assume the occurrence of some structural breaks in the time series. These could also contribute to the results of the unit root tests. To obtain further information about the occurrence of structural change in the data, we apply the Bai and Perron (2003) structural break tests, which allow us to obtain a number of interesting details about whether, when, and in which way structural breaks are present in the data. We assume that the time series values evolve in the following way:

$$x_t = \alpha_1 + \alpha_2 x_{t-1} + \varepsilon_t \tag{8}$$

If α_2 is equal to 1, this process characterizes a random walk plus drift, which can be assumed to be the data-generating process of many macroeconomic time series. The test identifies two breaks. The results for the three countries are displayed in Table A1.

In China, structural breaks are first present around the high tide of the Cultural Revolution (1966–1968). The second period during which breaks are found is the end of the Cultural Revolution in 1976 and the beginning of the first reform period (1979). Finally, the breaks in the early 1990s could be associated with Deng Xiaoping's Southern Tour (in 1992).

For Korea, we observe structural breaks during the high-growth phase from the mid-1960s to the mid-1970s. The beginning of the stabilization phase in 1980 appears to have had permanent effects on employment and output. The final period of structural breaks is related to the Asian Crisis.

For Japan, a relatively clear point to take away from the tests is that we observe structural breaks in the early 1970s. These are most likely related to the end of the fast-growth phase in 1972 and

[11] The unit root tests are not displayed but are available from the authors on request.

the beginning of the oil shock phase in 1973. The second break seems to occur towards the end of the bubble economy phase and at the beginning of the lost decade in the late 1980s. The employed persons' time series indicates instability during the Asian Crisis in 1997 but could also point towards the lost decade. The tests provide some evidence that the Japanese economy began to stagnate from about the late 1990s onwards, as the coefficient of α_2 is well below one.

3.5 Econometric framework

To determine whether the respective countries were following comparable growth paths towards high income, we use the concept of cointegration. Cointegration appears to be especially appropriate for our research question, as the notion of cointegration states that, between two or more non-stationary time series, a linear combination exists that generates residuals that are stationary (see Engle and Granger 1987). The error correction representation of the CVAR in the multivariate case is

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \phi D_t + \varepsilon_t \tag{9}$$

X_t is a vector that contains the variables included in the model. D_t is a vector of the deterministic components of the model. The Γ_i matrices contain the short-run information of the model, while Π contains the information about long-run relationships (which we are especially interested in) and can be rewritten as a vector product of $\alpha \cdot \beta'$. Here β' comprises the long-run information, while α contains the information on how and how quickly deviations from the long-run relations are corrected. Cointegration emerges when two or several non-stationary time series are driven by the same persistent shocks (Juselius 2006). In our case, we assume that the shocks driving our cointegrating relationships derive from economic policy measures implemented in the corresponding reform periods. Long-run relations can then be interpreted as economic steady-state relations. Extraordinary events can lead to outliers violating the normality assumption, as they lead to excess skewness and kurtosis (see Juselius 2018). Some of these problems can be resolved by including intervention dummies to account for significant political or institutional events. Less feasible seems to be the possibility to split the sample into more homogeneous periods, as only yearly data for our variables are available. A subsample analysis would hence suffer from problems associated with small samples.

Even if the residuals of the VAR pass the misspecification tests sufficiently well, this does not rule out the possibility that the model suffers from parameter non-constancy. For this reason, Juselius (2006) provides a description of several tests to assess the parameter constancy of the CVAR model. In the next section, we will discuss three selected tests for each model that we estimate to identify any remaining signs of structural change that we have not modelled.

First, we are especially interested in the stability of the long-run relationships, that is, the stability of $\hat{\beta}$. $\hat{\beta}$ can be seen as the average of the related coefficients over the sample period. It could be assumed that structural change induced by reforms affected not only the TFP growth but also the output elasticity of inputs. The optimal capital intensity could hence have changed (see Stijepic

and Wagner 2011). Theoretically this is also proved by the work of Acemoglu (2003). To test this, we perform a test of the 'known beta'. This test checks for the consistency of $\hat{\beta}$, that is, the stability of the long-run relations. The basic idea is that the model is estimated for a subsample period 1 to T_1, with $T_1 < T$, and then the recursive sample is extended until T is reached (see Juselius 2006).

Second, we apply recursive calculated prediction tests for the long-run relation. If the test value is above one, the model is not able to predict the observation for this period within the 95% confidence bands. This can be a helpful test to diagnose systematic predictive failure of the model (Juselius 2006).

Third, we can test the stability of the estimated coefficients over a sample period. This can give us additional insights into the stability of the long-run relations (Juselius 2006).

3.6 Estimation

The variable set that we use is relatively small. It is difficult to obtain data in sufficient quality that reach back far enough to observe signs of structural change. Multivariate cointegration analysis allows for multiple equilibrium relationships. Hence, with n variables in the system, up to n-1 cointegrating relations are possible. (Over-)identification of these relationships requires a sufficient amount of theoretical assumptions on what the concrete cointegrating relationships are. With a rising number of variables, it becomes more and more difficult to identify the 'true' structure of the data. In any case, cointegrating relationships that have already been identified should remain stable when additional variables are added. Hence, what we detect in the following should remain present in a richer information environment.

As mentioned in the previous section, we assume a simple Cobb–Douglas production function as the starting point of our analysis. Our vectors X_t for China, Korea, and Japan are

$$X_{j_t} = (Y_P_j_t, K_EMP_j_t, G_j_t, D_j_t) \tag{10}$$

for each country j.

We start by specifying the model for China. To test for cointegration, we first have to obtain a well-specified VAR model of the data. As TFP is usually captured by the residuals of the equation, we take into consideration the need to capture technological progress, structural change, and other variables that are not accounted for by the introduction of deterministic trends. Juselius (2006) generally proposes two approaches to identifying cointegrating relationships: the general-to-specific and the specific-to-general approaches. As we face a relative lack of restrictions and theories, we find it appropriate to start with a narrow model like that motivated by Chow (2015). Imposing restrictions on the model is crucial to decide how well the data fit our theoretical assumptions. We hence rearrange our baseline theoretical model to obtain a representation that allows us to impose over-identifying restrictions. Chow (2015) estimates several specifications of

his production function. The most specific is one with a fixed capital–labor ratio (k/l), which corresponds to capital intensity.

3.6.1 Analysis for China

3.6.1.1 Lag length selection and diagnostic testing of the unrestricted VAR

As we have yearly data, a lag length of order 1 appears to be appropriate, as it not very likely that information further back than 1 year is included in the investment and employment decisions. Lag reduction tests support this assumption. A lag length of 1 is superior to lag lengths of 2 or more. As the test statistics for the CVAR rely on the assumption of Gaussian residuals, deviation from the normality assumptions may distort the results. We hence must establish the necessary residual properties before we test for possible long-run relationships. From Sections 3.2 to 3.4, we already have a certain idea about possible outliers, trends, or (structural) breakpoints in the data. For China, we choose the sample period of 1969 to 2014. This excludes the 'abnormal' years of the Cultural Revolution (see Fig. 2 in Section 2). The Cultural Revolution was officially announced to be complete in 1969. We thus only miss out four values included by Chow (2015), whose sample is 1952–2012, omitting (1958–1969).[12] From our theoretical hypotheses in (7), we assume that we will have to include a deterministic trend in the model that allows for economic growth even if the input factors remain constant. We will, however, test the validity of this assumption. Chow (2015) argues that such a trend would catch increases in total factor productivity. Including a deterministic trend in the CVAR model is also the most general specification that one can choose. Hence, we include a deterministic trend over the whole sample period, additionally allowing for a break in this trend that mirrors a structural change induced by reforms or political events. Chow (1993) finds no evidence of TFP prior to 1980 and thus includes a broken linear trend beginning with $t = 1$ in 1979 (Chow 2015). With our data, a break in the deterministic trend appears in 1976 (often seen as the de facto end of the Cultural Revolution) such that $t = 1$ in 1977 seems to be more in line with the data. We additionally include intervention dummies for 1976, 1989, 1990, and 1991. We also include a shift dummy for 1990. With this specification, we obtain a relatively well-specified VAR model, as can be seen from the residual diagnostic tests in Table 2.

[12] Using a longer sample including earlier years leads to more dummy variables in our model and to a smaller, though significant, coefficient of the capital–employment ratio.

Table 2 Residual analysis – diagnostic testing of the unrestricted VAR (1) model: China

Multivariate test				
Residual autocorrelation:				
LM (1):	ChiSqr(4) = 5.742 [0.219]			
LM (2):	ChiSqr(4) = 0.879 [0.970]			
Test for ARCH:				
LM (1):	ChiSqr(9) = 11.875 [0.220]			
LM (2):	ChiSqr(18) = 16.430 [0.563]			
Univariate tests				
	ARCH(1)	Normality	Skewness	Kurtosis
Y_P_CHN	0.391 [0.532]	0.840 [0.657]	0.297	2.685
K_EMP_CHN	1.277 [0.258]	0.856 [0.652]	-0.160	3.050

Source: Own calculations.

3.6.1.2 Rank determination and testing restrictions on the CVAR

An LR test of long-run exclusion indicates that g_t is not part of the equilibrium relationship and hence we exclude it from the information set. Any external influences that are not accounted for by the variables in the system are caught by either the deterministic components or the residuals. The trace tests propose one cointegrating relationship, which we would also expect from the theory. We thus restrict the rank to be equal to one. This gives the following just-identified model (see Table 3).

Table 3: The just-identified long-run cointegration relations for r = 1: China

	Y_P_CHN	K_EMP_CHN	T(1976:01)	Trend
$\widehat{\beta}_1$	1.000	-0.453	-0.033	0.013
	(.NA)	(-7.189)	(-3.456)	(1.445)
$\widehat{\alpha}_1$	- 0.075	0.114		
	(-1.494)	(8.932)		

Source: Own calculations.

All the variables have the expected sign, and the gross domestic product per capita and the capital–employment ratio commove in a positive long-run relationship. TFP growth, proxied by the broken linear trend, has a positive effect on output and a significant coefficient of 0.033. In the next step, we relax the implicitly imposed restriction that the coefficients of capital and employment sum up to one by including capital and employment separately. By including the same dummy variables, we obtain a rather well-specified model without autocorrelation and mild skewness. The trace test indicates one cointegrating relationship. We continue to assume that the rank of Π is equal to one, which means that there is one single equilibrium relationship among the

variables. Again the trend is insignificant, as displayed in Table 3. We thus exclude the deterministic trend by imposing a 0 on the corresponding coefficient. Additionally, we explicitly test the coefficients of capital and labor to be of the same size. This over-identified model is displayed in Table 4.

Table 4 The over-identified long-run cointegration relations for r = 1: China

	Y_P_CHN	K_CHN	EMP_CHN	T(1976:01)	Trend
$\widehat{\beta}_1$	1.000	-0.417	0.417	-0.022	0.000
	(.NA)	(-7.278)	(7.278)	(-4.349	(.NA)
$\widehat{\alpha}_1$	-0.075	0.104	-0.031		
	(-1.534)	(8.079)	(-5.951)		

Source: Own calculations.

The over-identifying restrictions are accepted with a p-value of 0.135 and a $\chi^2(2)$ of 4.010.[13] The long-run equilibrium relationship is thus:

$$Y_P_{CHN_t} - 0.417 \cdot (K_{CHN_t} - EMP_{CHN_t}) \; - 0.022 \cdot \; trend76 \sim I(0) \tag{11}$$

Fig. 6 displays the corresponding cointegrating relationships. The residuals of the concentrated model (lower graph of Fig. 6) shows that the relationship is stable.

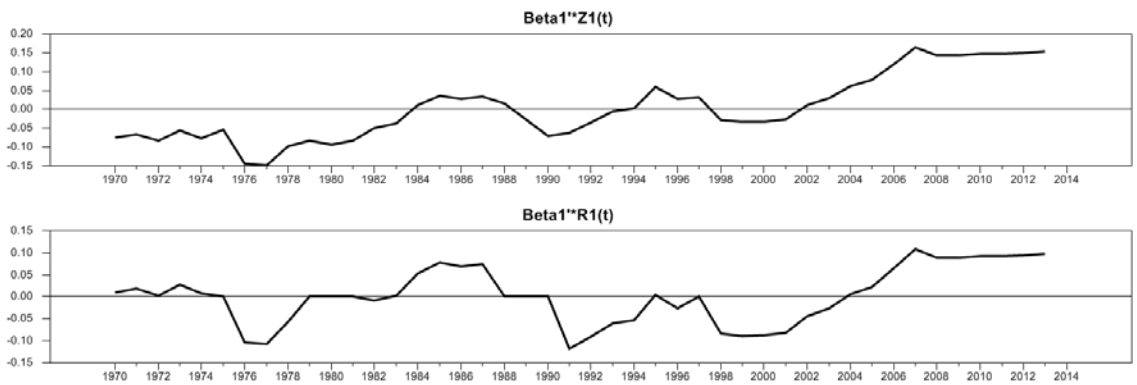

Fig. 6 Cointegrating relationships for China
Source: Own calculations.

3.6.1.3 Tests of constancy

[13] Juselius (2006) suggests applying the small-sample Bartlett correction in moderately sized samples of 50–70 observations, which applies in our case (45 observations). The non-corrected values are p = 0.063 and a chi-square (2) of 5.537 (correction factor: 1.381).

The test of beta constancy is displayed in Fig. 7. As we can see from Table A1, this is a very challenging test, as 1992 (the beginning of the subsample for the test in Fig. 7) is actually an important breakpoint in the Chinese data. The forward recursive calculated test shows that the model performs fairly well. After the assumed breakpoint in 1992 (Deng Xiaoping's Southern Tour of 1992), there is some instability, but the test statistic approaches its critical value of 1 relatively quickly. The Asian Crisis of 1997 and the Global Financial Crisis of 2007–2009 seem to have introduced some instability, but altogether we can assume that the coefficients remain stable after the Asian Crisis.

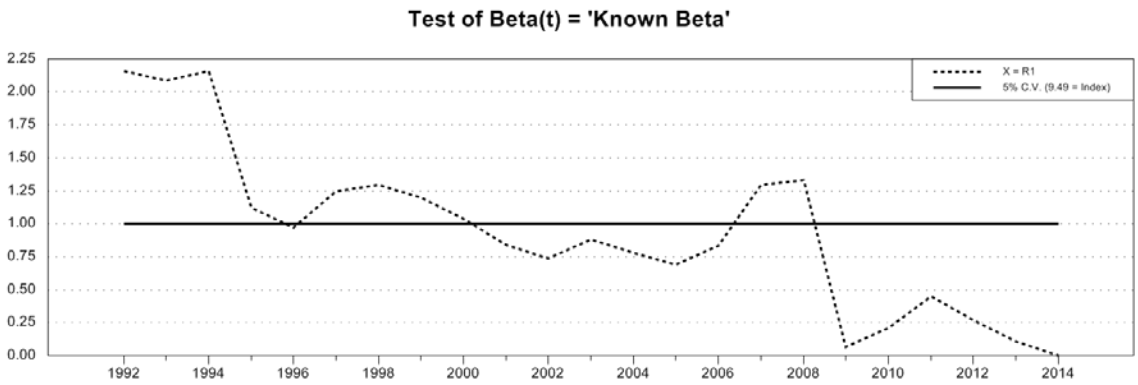

Fig. 7 Test of beta equal to the 'known beta': China
Source: Own representation.

The one-step prediction tests in Fig. 8 show that major predictions errors can again be related to Deng Xiaoping's Southern Tour of 1992, a major reform in 1994 (see Cheremukhin et al. 2015), and the Asian and the Global Financial Crisis.

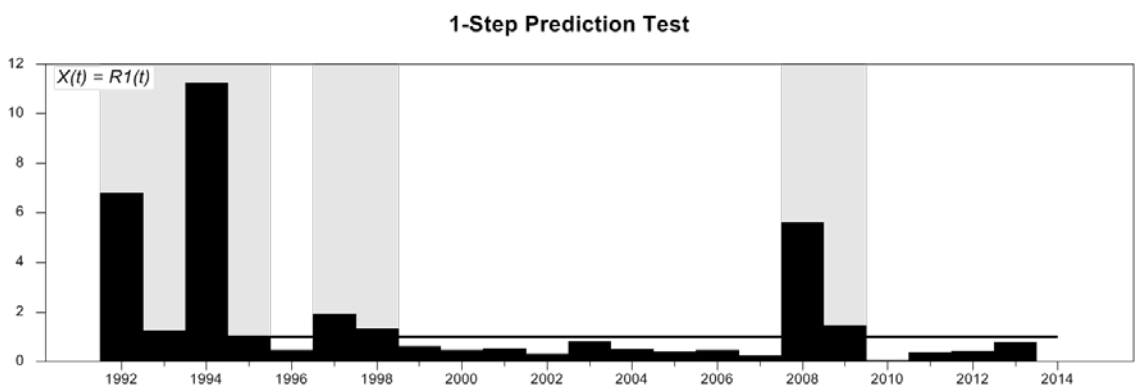

Fig. 8 One-step prediction test for the concentrated model
Source: Own calculations.

As can be seen in Fig. 9, the coefficients of capital and labor remain relatively stable. We cannot deny that there seems to be a visible shift in all the coefficients roughly after the end of the Asian

Crisis in 1998, as the coefficient of capital intensity is somewhat higher for the later part of the sample. There are two explanations for this fact. Following the arguments of Chow (2015), a higher coefficient for capital could be explained by the still relative abundance of labor in China. Another explanation could be that capital was actually invested more efficiently over time and hence received a larger share of production. TFP growth also seems to have profited from China's WTO entry in 2001. However, this is somewhat speculative, as all three coefficients only show minor fluctuations.

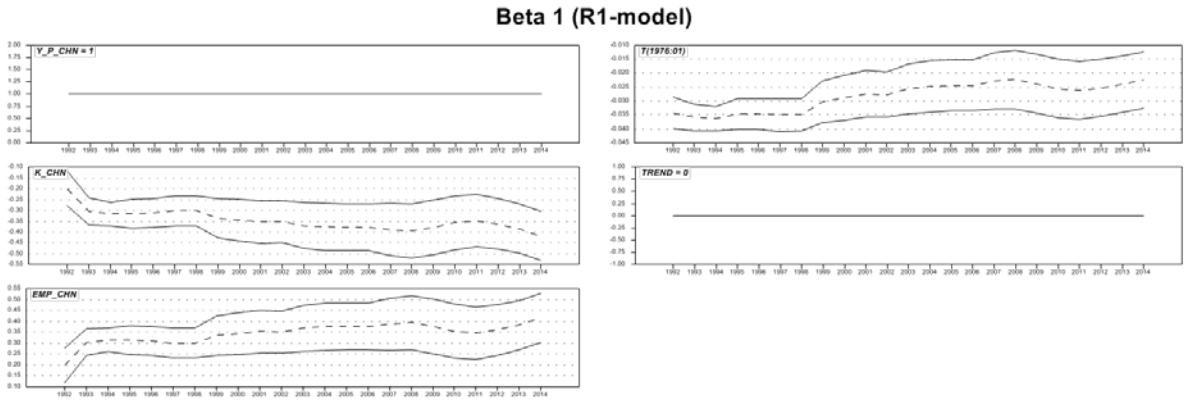

Fig. 9 Test of coefficient constancy: China
Source: Own calculations.

We find that the long-run evolution of per capita GDP, capital stock, and employment in China can be reconciled with Cobb–Douglas production functions, which is the first notable result. Coefficients of capital stock and labor that sum up to 1 are not rejected. We find evidence of a (broken) linear trend in the cointegrating relations that can point to changes in TFP growth. The structure of the data is reconcilable with the hypothesis that TFP growth in China started around 1977.

For China, our model in (11) is able to reproduce the main features of Chow (2015). We estimate a coefficient of 0.41 for the capital share. This coefficient is smaller than Chow's estimates but relatively close. Chow (1993) also estimates production functions for non-agricultural sectors. The coefficients are highest for the industry sector but much smaller for service-related sectors, like commerce and transportation.[14] Given that our sample includes 2 more years in which the service sector share in employment in China has increased strongly,[15] this result seems to be acceptable.

However, the choice of sample has a strong impact on the coefficient size, as the inclusion of some of the abnormal years leads directly to a lower coefficient for the capital share. Our coefficient for TFP growth of 0.022 is similar in size to the estimates in Chow (2015), which

[14] See Chow (1993, p. 833, Table XII).
[15] See Murach and Wagner (2017, p. 273, Fig. 1).

range between 0.025 and 0.030.[16] In addition, our use of Y/N instead of Y or Y/L in our estimations may lead to lower estimates. As pointed out by Chow (2015), a higher coefficient for capital points towards relative abundance of labor relative to capital. We could hence also argue that, although still considerable in an international comparison, the reservoir of labor is slowly dropping relative to earlier periods.

To obtain (11), we make extensive use of the ex-ante information of Section 3.2 and Section 3.4. We exclude most of the abnormal years of the 1950s and 1960s, and our sample thus starts in 1969. This year appears to signal a structural break (see Table A1) that could be an indication of the economic take-off, as the point of fertility decline in China is reported as 1970. TFP growth from 1977 onwards could also be justified with the results from Table A1, which indicates a structural break between the years 1976 and 1979. The one-step prediction test in Fig. 8 shows that the breakpoint in 1991/1992 (possibly related to the Southern Tour) is (only) of a transitory nature. We hence feel that we have convincingly modelled the major features of the Chinese economy. The fact that the cointegration relation for the concentrated model in Fig. 6 increases between 2001 and 2007 could be interpreted as the positive effect of China's WTO entry in 2001. However, this does not become a break in TFP, because the financial crisis halted it. We are a little disappointed with the estimated alpha coefficients of the long-run relationships. Table 4 indicates that capital and employment show error-correcting behavior to deviations from the long-run relationship, but the per capita GDP appears to be weakly exogenous.

3.6.2 Korea

3.6.2.1 Lag length selection and diagnostic testing on the unrestricted VAR

For the Korea model, the lag reduction tests point to a lag length of 1. There is some controversy about the sources of growth in Korea. For instance, Young (1995) argues that most of the dynamic growth in the so-called Four Asian Tigers can be explained by factor accumulation and labor reallocation between sectors. Accounting for these explanatory variables would lead to much lower estimates of the total factor productivity growth. In a more recent analysis by Jeong (2017), the results are similar, as it is explained that human capital was the main driver of growth during the 1960s and capital deepening in the 1970s. Productivity growth was then the main driver of growth in the 1980s, 1990s, and 2000s. This would justify two possible specifications of deterministic components. Either we could include a deterministic trend over the whole sample period or we could include such a trend only from 1980 onwards with the initiation of the stabilizing stage in Korea.

We start with the most general setting concerning deterministic components in our model. We allow the data to be trend-stationary and to have non-zero intercepts to see whether trends persist or disappear in the cointegrating relations. We first analyze the specific relation between the gross domestic product per capita and the capital–employment ratio. In this model, we decide to

[16] See Chow (2015, p. 100, Table 5.2); values rounded to three decimals.

use four dummy variables for the years 1968, 1969, 1980, and 1998. The diagnostic tests are presented in Table 5. The model can be regarded as being well specified.

Table 5 Residual analysis – diagnostic testing of the unrestricted VAR (1) model: Korea

Multivariate test				
Residual autocorrelation:				
LM (1):	ChiSqr(4) = 7.570 [0.109]			
LM (2):	ChiSqr(4) = 3.874 [0.423]			
Test for ARCH:				
LM (1):	ChiSqr(9) = 7.939 [0.540]			
LM (2):	ChiSqr(18) = 9.929 [0.934]			
Univariate tests				
	ARCH(1)	Normality	Skewness	Kurtosis
Y_P_KOR	0.043 [0.835]	0.480 [0.787]	0.056	2.346
K_EMP_KOR	1.627 [0.202]	2.003 [0.367]	-0.380	3.304

Source: Own calculations.

3.6.2.2 Rank determination and testing restrictions on the CVAR

The trace test indicates one cointegrating relationship, as displayed in Table 6.

Table 6 LR trace test for the unrestricted VAR (1) model: Korea

r	p-r	Eigenvalue	Trace	95% crit. value	p-value
2	0	0.792	95.414	36.970	0.000
1	1	0.232	13.704	18.547	0.218

Source: Own calculations.

Normalizing on the per capita GDP, we obtain the following just-identified model (see Table 7).

Table 7 The just-identified long-run cointegration relations for r = 1: Korea

	Y_P_KOR	K_EMP_KOR	T(1980:01)	Trend
$\widehat{\beta}_1$	1.000	-0.560	-0.032	-0.001
	(.NA)	(-6.378)	(-8.634)	(-0.112)
$\widehat{\alpha}_1$	0.130	0.221		
	(5.604)	(13.091)		

Source: Own calculations.

Assuming that TFP growth set in around 1980 supports the assumption that the forces that drove Korea's growth actually changed during this period. Introducing a broken linear trend in 1980 actually renders the TFP growth over the whole period (close) to zero and more importantly insignificant (while still having the correct sign).

In the next step, we loosen our restrictions by allowing capital and employment to move independently. We assume that TFP growth only set in significantly around 1980. With the same dummy variables as in the previous modelling cycle, we obtain a fairly well-specified model of the data. Table 8 displays the diagnostic test of the residuals of the VAR(1) with three variables.

Table 8 The over-identified long-run cointegration relations for r = 1: Korea

	Y_P_KOR	K_KOR	EMP_KOR	T(1980:01)	Trend
$\hat{\beta}_1$	1.000	-0.667	0.667	-0.022	0.000
	(.NA)	(-36.866)	(36.866)	(-17.461)	(.NA)
$\hat{\alpha}_1$	0.160	0.326	0.031		
	(4.973)	(20.176)	(1.789)		

Source: Own calculations.

The over-identifying restrictions are accepted with a p-value of 0.324 and a $\chi^2(2)$ of 0.324.[17] The long-run equilibrium relationship for Korea is thus:

$$Y_P_{KOR_t} - 0.667 \cdot (K_{KOR_t} - EMP_{KOR_t}) - 0.022 \cdot trend80 \sim I(0) \qquad (12)$$

Fig. 10 displays the corresponding cointegration graph.

[17] This is the Bartlett corrected value.

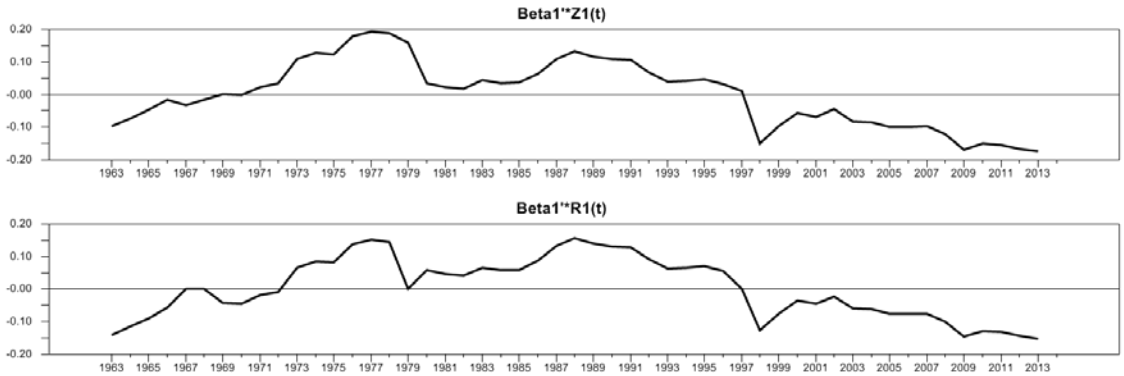

Fig. 10 Cointegrating relations for Korea
Source: Own calculations.

The residuals of the concentrated model appear to be fairly stationary. A visual inspection cannot rule out the possibility that some TFP growth was also present prior to the early 1980s.

3.6.2.3 Tests of constancy

The test of beta constancy in Fig. 11 shows that there was some instability in the long-run relations at the beginning of the stabilization stage in the 1980s. However, the test statistics remain under the critical value of one and hence the beta coefficients were stable over this subsample period.

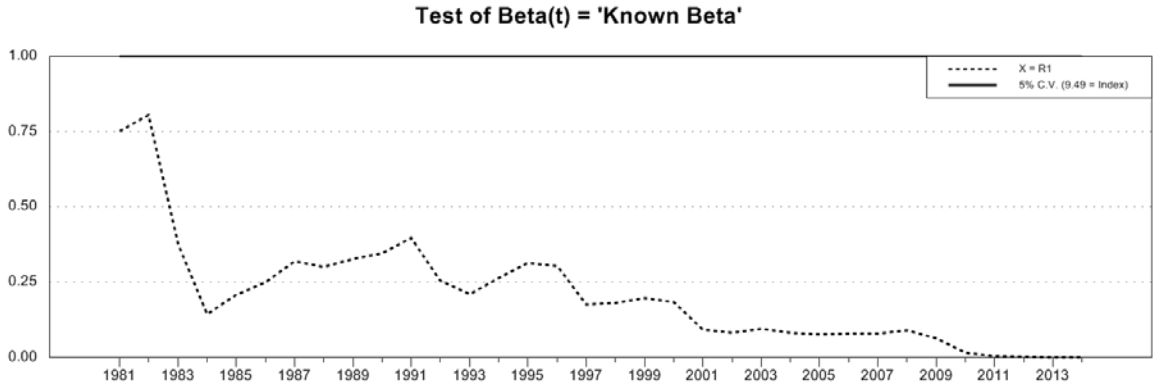

Fig. 11 Test of beta equal to the ‘known beta’: Korea
Source: Own representation.

The one-step prediction test (see Fig.12) for the concentrated model performs quite well. At the beginning of the stabilizing stage in the early 1980s, it shows a few noticeable prediction errors. The Asian Crisis in 1997 generated a single and very high prediction error. For the remaining years, the model is able to predict the developments very well.

Fig. 12 One-step prediction test for the concentrated model: Korea
Source: Own calculations.

The coefficients displayed in Fig. 13 of the long-run relations in the case of Korea remained relatively stable over the subsample period. There was some instability from around 1987 until 1990.

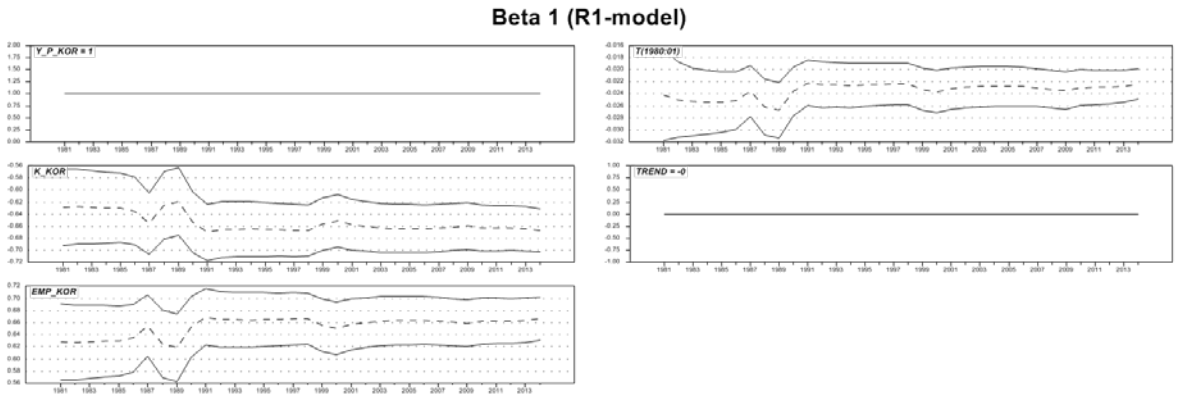

Fig. 13 Test of coefficient constancy: China
Source: Own calculations.

3.6.3 Analysis for Japan

3.6.3.1 Lag length selection and diagnostic testing of the unrestricted VAR

For Japan, the lag reduction tests point towards a lag order of 1. Additionally to the deterministic trend, we include one breakpoint ($t = 1$ in 1973) to account for changes in the light of the oil crisis. Additionally, we include dummies for 1973 and 1997 (shift dummies) as well as an intervention dummy for 2009. We thus obtain a fairly well-specified model, as displayed in Table 9.

Table 9 Residual analysis – diagnostic testing of the unrestricted VAR (1) model

Multivariate test				
Residual autocorrelation:				
LM (1):	ChiSqr(9) = 31.703 [0.000]			
LM (2):	ChiSqr(9) = 20.288 [0.016]			
Test for ARCH:				
LM (1):	ChiSqr(36) = 52.761 [0.035]			
LM (2):	ChiSqr(72) = 84.896 [0.142]			
Univariate tests				
	ARCH(1)	Normality	Skewness	Kurtosis
Y_P_JP	1.583 [0.208]	1.007 [0.604]	-0.048	3.157
K_EMP_JP	0.245 [0.621]	1.057 [0.589]	-0.021	3.173
DG_JP	0.127 [0.721]	1.162 [0.559]	-0.261	3.107

Source: Own calculations.

3.6.3.2 Rank determination and testing restrictions on the CVAR

The rank test indicates one or two cointegrating relationships (see Table 10). Additionally, this time, the test of long-run exclusion does not suggest that g_t is not part of the cointegrating relationships. We assume a rank of 1.

Table 10 LR trace test for the unrestricted VAR (1) model

r	p-r	Eigenvalue	Trace	95% crit. value	p-value
3	0	0.828	143.722	58.737	0.000
2	1	0.431	41.084	37.115	0.015
1	2	0.144	8.941	18.800	0.631

Source: Own calculations.

This model gives the long-run relationship displayed in Table 11.

Table 11 The just-identified long-run cointegration relations for r = 1

	Y_P_JP	K_EMP_JP	G_JP	T(1972:01)	Trend
$\hat{\beta}_1$	1.000	-0.283	0.829	0.037	-0.046
	(.NA)	(-17.209)	(4.245)	(22.066)	(-23.745)
$\hat{\alpha}_1$	-0.004	0.422	0.044		
	(-0.093)	(15.554)	(2.694)		

Source: Own calculations.

The capital–employment coefficient is comparably small. The trend points to overall TFP growth over the whole sample. The break in 1973 sets off most of it. Afterwards, TFP growth would only be as high as 0.08 on average. We also note that g_t is significant and very close to the theoretical value of 1 that we derived in Section 3.1.

In the model including capital and employment separately, the inclusion of g_t is not rejected as well. Here, the trace test indicates two cointegrating relationships, and we try to impose the theoretical relation developed in (7) on the second relation. This gives the following over-identified model (see Table 12).

Table 12 The over-identified long-run cointegration relations for r = 1: Japan

	Y_P_JP	K_JP	EMP_JP	G_JP	T(1972:01)	Trend
$\widehat{\beta}_1$	0.000	-0.048	1.000	0.832	0.009	-0.008
	(.NA)	(-22.433)	(.NA)	(29.930)	(40.292)	(-27.272)
$\widehat{\beta}_2$	1.000	-0.264	0.264	1.000	0.044	-0.051
	(.NA)	(-23.069)	(23.069)	(.NA)	(36.361)	(-33.139)

Source: Own calculations.

These restrictions are accepted with a p-value of 0.493 and a CHISQR (1) of 0.493. The second cointegrating relationship is our estimate for the production function.

$$Y_P_{JP_t} - 0.264 \cdot (K_{JP_t} - EMP_{JP_t}) - 1.000 \cdot G_{JP_t} - 0.051 \cdot trend + 0.044 \cdot trend72 \sim I(0) \tag{13}$$

Fig. 14 displays the corresponding cointegration graph.

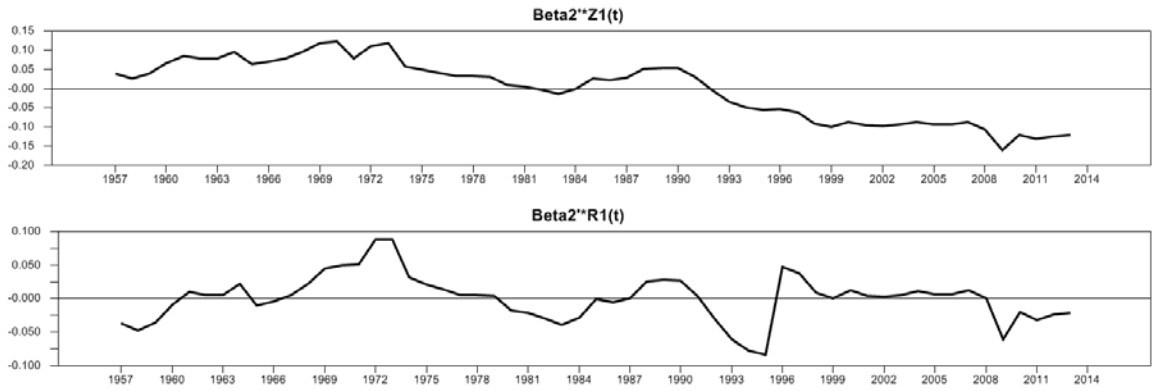

Fig. 14 Second Cointegrating relations for Japan
Source: Own calculations.

3.6.3.3 Tests of constancy

The test of beta constancy indicates extended periods of instability of the beta coefficients from the mid-1980s onwards. Towards the Asian Crisis, instability decreased considerably but remained significant. The revitalization phase fits with the overall stable beta coefficients (see Fig. 15).

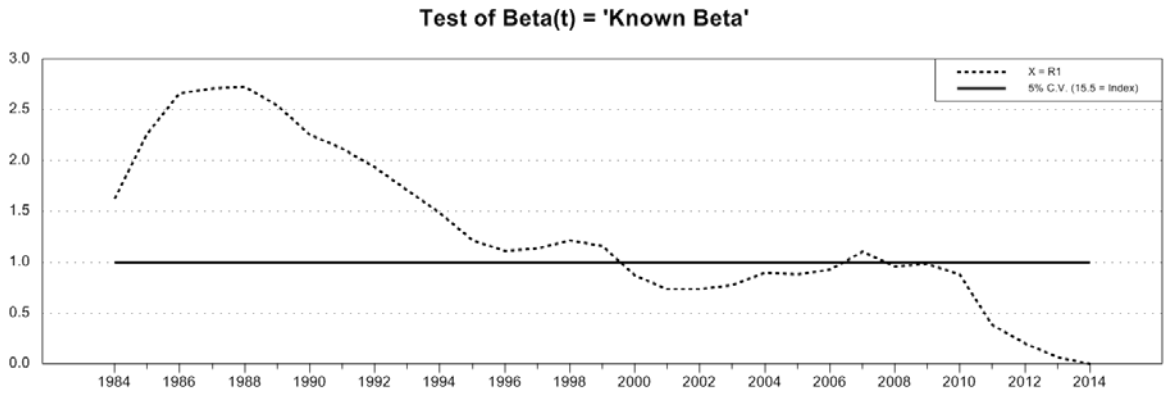

Fig. 15 Test of beta equal to the ‘known beta’: Japan
Source: Own representation.

The one-step prediction tests in Fig. 16 give the impression that the model fails to make trustworthy predictions for the sample period. While the two largest prediction errors are related to the Asian Crisis and the Global Financial Crisis, the model also has increasing difficulties in predicting the developments during the bubble economy phase of 1985 to 1990 and the beginning of the lost decade.

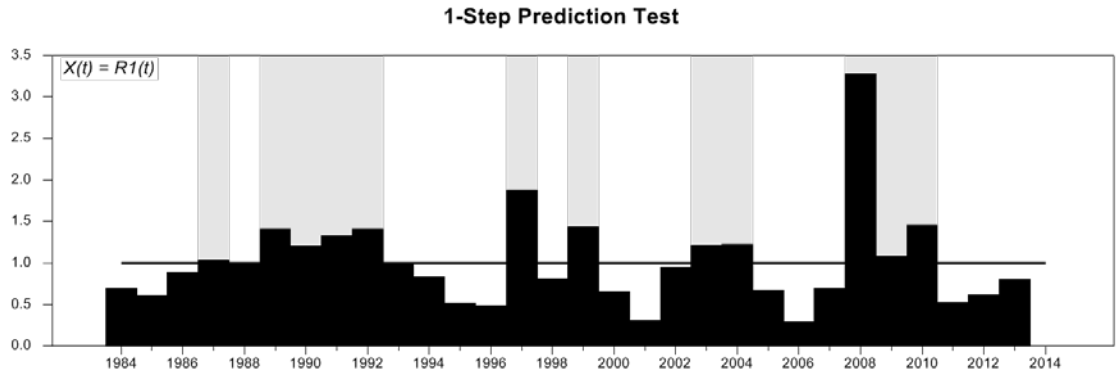

Fig. 16 One-step prediction test for the concentrated model: Japan
Source: Own calculations.

The coefficients displayed in Fig. 17 of the long-run relations in the case of Japan remained very unstable over the phase of the bubble economy in the 1980s. They give a clear indication that the Japanese economy was not on its equilibrium path during this period.

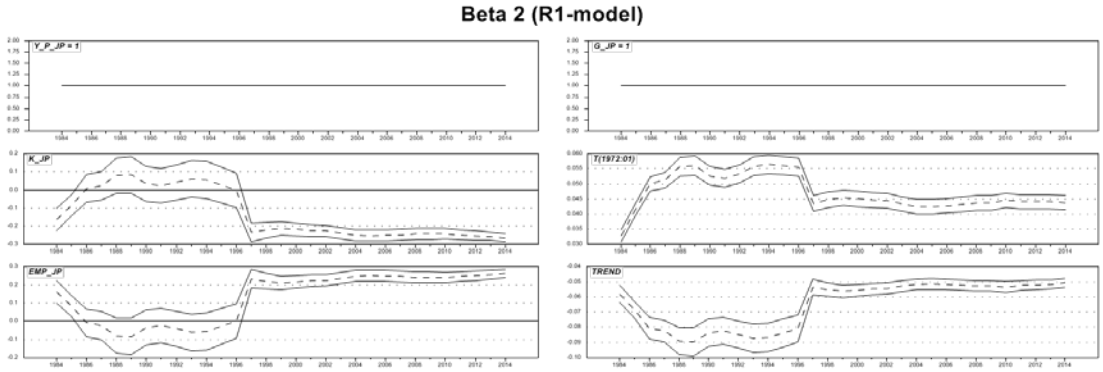

Fig. 17 Test of coefficient constancy: Japan
Source: Own calculations.

4 Conclusion

Sustained rapid growth has transformed China from a low-income economy into a middle-income economy in a remarkably short period of time. The next challenge for China is to graduate from middle income to high income. How well and smoothly China tackles this difficult challenge has sizable ramifications not only for China but, given China's large and growing footprint on the global economy, for the rest of the world.

In this paper, we sought to obtain some clues about the future dynamics of China's economic growth by looking at Japan's, Korea's, and China's past pattern of growth. Accordingly, we analysed and compared the structural change and growth experiences of the three countries,

which share many similarities. Section 2 pointed out these similarities. Perhaps the most significant common denominator was capable bureaucracy and a developmental state that prioritized economic growth (favoring export- and investment-led growth) and played the role of a catalyst in the rapid growth and structural transformation of the three countries. As such, structural policies and reform played a major role.

Our analysis and comparison of the patterns of economic growth and structural change in Japan, Korea, and China yielded a number of interesting findings. The pattern of structural change in Japan, Korea, and China seems to have been broadly similar, with China following in Korea's footsteps by about 25 to 30 years and following Japan by 50 to 55 years. The descriptive analysis of Section 2 and the more in-depth econometric analysis of Section 3 both support the view that many features of China's economic development mirror the earlier Japanese and Korean experience. The GDP growth and capital–labor ratio moved together in a positive long-run relationship in all countries, which can be brought into line with the hypotheses derived from Section 3.1. An interesting finding is hence that the export- and investment-led growth models that all three countries followed for an extensive period of time are reconcilable with Cobb–Douglas production functions with broken linear trends.

However, there is one interesting and significant difference between the three countries in their economic growth and structural change trajectory. Specifically, our analysis indicates that China experienced growth based on total factor productivity (TFP) gains at a much earlier stage in its development path than Korea. If China shifted towards TFP growth at a similar stage to Korea, the shift would have occurred around 2011. In fact, the shift in China began as early as the late 1970s. In comparison with Japan, there are no visible signs that TFP growth in China slowed down significantly in our sample period.

The broader question that we sought to address through our comparative analysis of the growth and structural change experiences of China, Japan, and Korea is whether China can replicate especially Korea's success in graduating smoothly from middle income to high income in a relatively short period of time. At a broader level, the balance of evidence from our analysis provides cautious grounds for optimism about China's prospects for a smooth and quick transition to high income. Above all, the fact that China's growth has been led by TFP growth in addition to factor accumulation suggests that it may be sustainable.

However, to continue its enviable track record of rapid TFP growth, China must forcefully implement structural reforms, such as state-owned enterprise (SOE) reform and reducing the role of the state in the financial system. Structural challenges, such as population aging, and new risks, such as rising global protectionism, further strengthen the case for such TFP-promoting reforms.

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Appendix

Table A1 Structural break tests

			Japan			
Y_P_JP	Lower 95%	Upper 95%		**≤ 1970**	**≤ 1987**	**> 1987**
1970	1969	1972	α_1	-0.103	0.134	1.956
1987	1976	1988	α_2	1.021	0.990	0.813
K_JP	Lower 95%	Upper 95%		**≤ 1972**	**≤ 1988**	**> 1988**
1972	1971	1973	α_1	-0.333	0.764	2.035
1988	1987	1989	α_2	1.035	0.956	0.878
EMP_JP	Lower 95%	Upper 95%		**≤ 1988**	**≤ 1997**	**> 1997**
1988	1984	1989	α_1	0.155	0.843	1.546
1997	1991	1998	α_2	0.964	0.800	0.629
K_EMP_JP	Lower 95%	Upper 95%		**≤ 1972**	**≤ 1988**	**> 1988**
1972	1971	1973	α_1	-0.303	0.602	1.350
1988	1987	1989	α_2	1.046	0.954	0.893
G_JP	Lower 95%	Upper 95%		**≤ 1988**	**≤ 1997**	**> 1997**
1988	1986	1991	α_1	0.116	0.171	0.220
1997	1994	1998	α_2	0.823	0.723	0.674
			Korea			
Y_P_KOR	Lower 95%	Upper 95%		**≤ 1967**	**≤ 1982**	**> 1982**
1967	1965	1969	α_1	-1.103	0.411	0.487
1982	1981	1986	α_2	1.151	0.958	0.956
K_KOR	Lower 95%	Upper 95%		**≤ 1969**	**≤ 1997**	**> 1997**
1969	1968	1970	α_1	-1.103	0.411	0.487
1997	1996	2008	α_2	1.151	0.958	0.956
EMP_KOR	Lower 95%	Upper 95%		**≤ 1975**	**≤ 1984**	**> 1984**
1975	1974	1976	α_1	-0.079	0.639	0.196
1984	1983	1985	α_2	1.052	0.761	0.942
K_EMP_KOR	Lower 95%	Upper 95%		**≤ 1965**	**≤ 1998**	**> 1998**
1965	1964	1966	α_1	0.639	0.097	0.646
1998	1997	2002	α_2	0.930	0.999	0.950
K_EMP_KOR	Lower 95%	Upper 95%		**≤ 1964**	**≤ 1984**	**> 1984**
1964	1962	1965	α_1	0.214	0.052	0.051
1984	1977	1989	α_2	0.838	0.943	0.921
			China			
Y_P_CHN	Lower 95%	Upper 95%		**≤ 1968**	**≤ 1976**	**> 1976**
1968	1968	1990	α_1	3.645	3.103	-0.046
1976	1975	1978	α_2	0.473	0.564	1.013
K_CHN	Lower 95%	Upper 95%		**≤ 1969**	**≤ 1992**	**> 1992**
1969	1968	1982	α_1	0.966	0.090	-0.054
1992	1990	2005	α_2	0.933	0.999	1.010
EMP_CHN	Lower 95%	Upper 95%		**≤ 1963**	**≤ 1979**	**> 1979**
1963	1963	1970	α_1	1.394	0.121	0.430
1979	1977	1980	α_2	0.751	0.985	0.936
K_EMP_CHN	Lower 95%	Upper 95%		**≤ 1969**	**≤ 1991**	**> 1991**
1969	1968	1978	α_1	0.989	-0.031	-0.061
1991	1990	1996	α_2	0.881	1.008	1.015
G_CHN	Lower 95%	Upper 95%		**≤ 1964**	**≤ 1982**	**> 1982**
1964	1963	1979	α_1	0.258	-0.080	0.035
1982	1981	1983	α_2	0.712	1.081	0.933

Source: Own calculations (sample: 1956–2014).

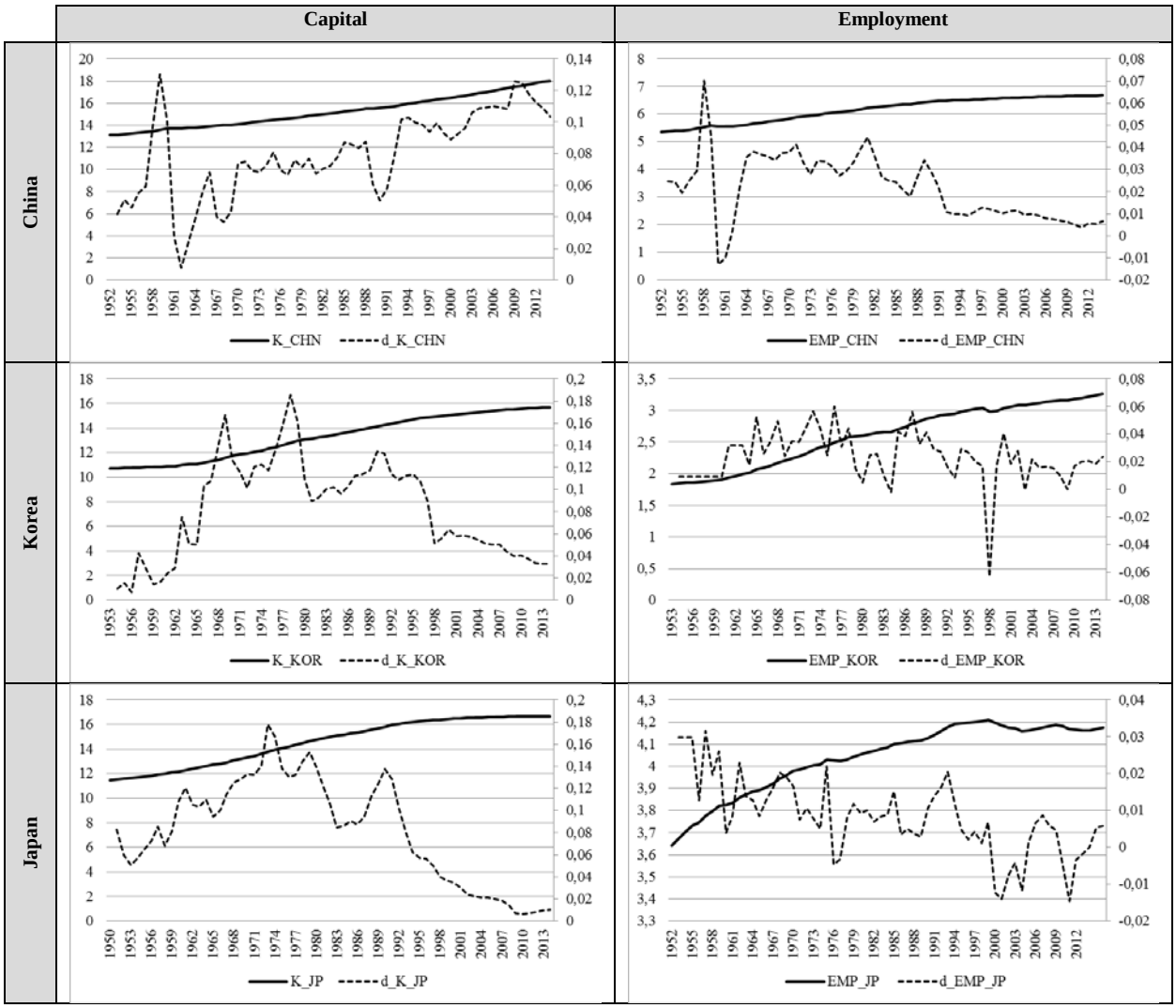

Fig. A1 Time series of capital and employment
Source: Own calculations.

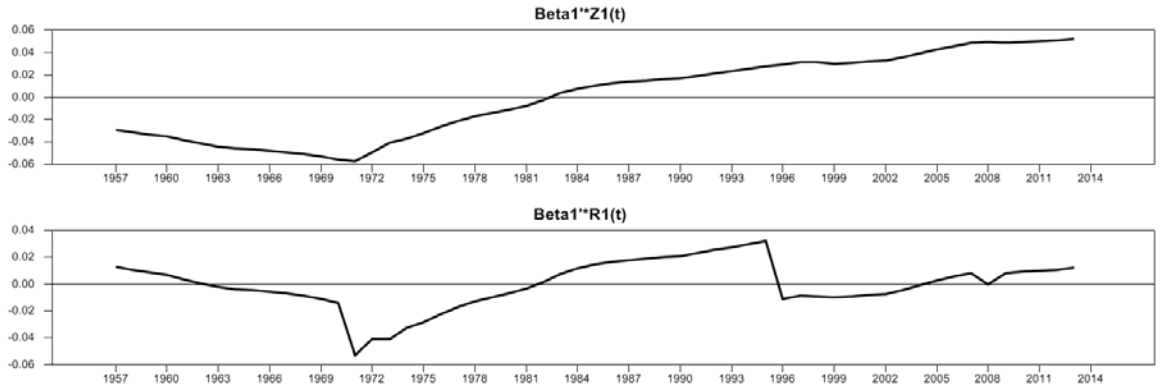

Fig. A2 First cointegration relations: Japan
Source: Own calculations

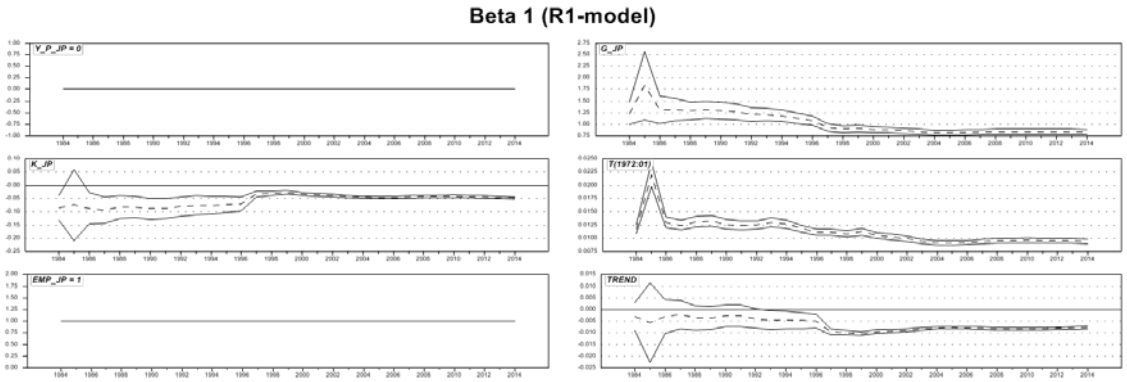

Fig. A3 Coefficient stability for the first cointegration relation: Japan
Source: Own calculations.