Chukyo University Institute of Economics

Discussion Paper Series

August 2019

No. 1904

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Kazuyuki Inagaki
Faculty of Economics, Nanzan University

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Kazuyuki Inagaki
Faculty of Economics, Nanzan University
18 Yamazato-cho, Showa-ku, Nagoya 466-8673 JAPAN
E-mail address: inagaki@nanzan-u.ac.jp

ABSTRACT

Using a nonlinear cointegration technique, this paper shows that the bilateral U.S. current account balance with China has a U-shaped relationship with the life expectancy gap between the United States and China. A narrowing gap means that the catch-up in Chinese life expectancy initially increases the U.S. deficit. However, the direction of this effect is reversed as the life expectancy gap continues to narrow, so that the increased U.S. deficit with China falls with a further catch-up in Chinese life expectancy. The life expectancy gap has been below its threshold level since 2013, and the United Nations predicts that the gap will fall over the next 60 years. This long-run demographic trend has the potential to improve the U.S. current account balance with China. To provide a theoretical foundation for the empirical results, we develop a simple two-country overlapping generations model and show that the U-shaped relationship can be theoretically reproduced.

Keywords: U.S. current account balance with China; Nonlinear cointegration; Catch-up in life expectancy

JEL Classification Codes: F32; J11

Acknowledgements

I am highly grateful to Martin Wagner for his kind help. This work was supported by JSPS KAKENHI Grant Number 16K17147, and I have no conflict of interest. I am solely responsible for any errors.

1. Introduction

This paper addresses the question: "how is the United States affected by the catch-up of China's life expectancy?" This question is motivated by the fact that the convergence speed of life expectancy at birth is much higher in China than in other middle-income countries (Figure 1). As Krueger and Ludwig (2007) suggest, a rapid increase in the life expectancy of a trading partner affects capital export to a home country, and therefore, the catch-up of China's life expectancy is likely to have a significant impact on the U.S. external deficit.[1] However, the nature of this impact remains unclear. In the literature on aging and health in China, many studies pay attention to domestic issues such as economic growth and social security reforms (e.g., Ebenstein, et al., 2015; Li and Lin, 2016). We extend this literature in the framework of open-economy macroeconomics.

Figure 1 around here

To assess the international spillover effect of the rapid increase in Chinese life expectancy, this paper focuses on the bilateral U.S. current account balance with China (i.e., net capital flows from China to the United States). Its ratio to the total U.S. current account balance increased from 25% in 2003 to 77% in 2016, and the U.S. deficit with China is closely related to the U.S. net international investment position (i.e., the difference between external financial assets and liabilities) that exceeds negative 30% of U.S. GDP (Ahmed et al., 2018). The size of the U.S. deficit with China is also a major concern for the world economy because their trade tensions are expected to increase the risk of an abrupt global slowdown (World Bank, 2019). However, surprisingly little attention has been paid to the determinants of the bilateral U.S. current account balance with China. Although data on this variable are available from 2003, this period includes sufficient information following China's accession to the WTO and is suitable for analyzing cross-border effects of Chinese capital as mentioned above. The U.S. external deficit and population aging in China are timely and important issues, and their connection is of great interest.

Our empirical analysis is in line with the literature on the determinants of current account balances. It is well accepted that population aging affects current account balances,

[1] Life expectancy is associated with the life-cycle behavior of individuals including their saving decision. From a macroeconomic perspective, the resulting change in capital accumulation affects capital flows, and this effect is measured by current account balances (i.e., domestic savings minus domestic investment).

and the old-age dependency ratio is a commonly used measure (e.g., Chinn and Prasad, 2003; De Santis and Lührmann, 2009; Higgins, 1998; Kim and Lee, 2008). We extend the previous studies by decomposing the old-age dependency ratio into life expectancy and the growth rate of working-age population.[2] Our new finding is that the former's effect exhibits a significant sign change in the process of catch-up in life expectancy.

Specifically, the life expectancy gap (defined as U.S. life expectancy minus Chinese life expectancy) is used to measure the catch-up in life expectancy. We find that this variable has a U-shaped impact on the U.S. current account balance with China as described in Figure 2. Therefore, a narrowing life expectancy gap initially increases the U.S. deficit with China, but eventually the increased U.S. deficit falls with the further catch-up of Chinese life expectancy. This result is confirmed after controlling for major determinants of current account balances including business cycles, fiscal balances, international competitiveness, financial development, and financial crises (e.g., Arghyrou and Chortareas, 2008; Chinn and Prasad, 2003; Chinn and Ito, 2007; De Santis and Lührmann, 2009; Gruber and Kamin, 2007; Lane and Milesi-Ferretti, 2012; Unger, 2017). Furthermore, we report 64 estimation results in total and the U-shaped relationship is robust for all cases.

Figure 2 around here

This paper also makes econometric contributions by using a recently developed technique of polynomial cointegration (Wagner, 2015; Wagner and Hong, 2016) for the first time in the literature. The U-shaped relationship in Figure 2 is demonstrated by this technique. As Feroli (2006) emphasizes, the persistent trend of population aging has a long-run impact on current account balances, and regression models consisting of cointegrated variables can be regarded as long-run equilibrium relationships (Engle and Granger, 1987). Many studies use linear cointegrating regression models to examine the long-run determinants of current account balances for various countries (e.g., Arghyrou and Chortareas, 2008; Belke and Dreger, 2013; Hung and Gamber, 2010; Unger, 2017). Along this line, our long-run equation includes commonly used variables in linear form, but only the life expectancy gap is allowed to have a quadratic polynomial relationship to the U.S. current account balance with China. This small nonlinear adjustment improves the out-of-sample forecast accuracy of our model by 74%

[2] This decomposition is justified theoretically. A detailed explanation is provided in Section 6.1.

relative to the standard linear specification.

Our finding for the U-shaped impact differs somewhat from the hypothesis provided by the existing empirical literature, which suggests that foreign life expectancy has a linear positive impact on the current account balance in the home country.[3] To provide a theoretical foundation for our empirical results, we develop an overlapping generations (OLG) model in an open-economy setting, which is simple and analytically tractable and thus serves to provide a clear understanding of the mechanism of the U-shaped impact. Specifically, we derive the general equilibrium effect of the rise of the foreign life expectancy as follows.

(i) Young agents in the foreign country increase savings to finance consumption in a longer retirement period, and this effect promotes capital flows to the home country.

(ii) The social security burden increases as the number of old agents increases, which reduces foreign savings and thus capital flows to the home country.

(iii) Better health improves productivity in old age.[4] In the foreign country, individuals who expect to earn more in old age save less over their working life, and this decision reduces capital flows to the home country.

Our comparative static analysis indicates that effect (i) is dominant at the early stage of the catch-up process of foreign life expectancy. However, the sum of effects (ii) and (iii) exceeds effect (i) when foreign life expectancy increases further. Therefore, the current account deficit in the home country increases at first but eventually decreases in the process of life expectancy convergence between the home and foreign countries. An important fact is that the savings rate in China has increased for much of the past 20 years but started to decrease in recent years (see Section 5.2). This fact is consistent with our theoretical results explained above.

Effects (i) and (ii) are consistent with the results of Ito and Tabata (2010), and we extend their analysis by incorporating the interaction effect between life expectancy and elderly labor supply (iii) into our theoretical model. This extension is based on the fact that the labor force participation rate of the elderly has increased remarkably in countries with rapidly aging

[3] In short, because old agents draw down their savings, an increase in the foreign life expectancy (or the resulting increase in the foreign old-age dependency ratio) worsens the foreign current account balance. This leads to the same amount of improvement in the home current account balance.

[4] This assumption is based on the model developed by Aísa et al. (2012) and is plausible because higher life expectancy means better health. We find that this assumption holds empirically. The discussion and empirical results are reported in Section 6.2 and Appendix B.

populations.[5] In addition, we show that, under the parameter values calculated from U.S. and Chinese data, the interaction effect (iii) is essential for reproducing the U-shaped impact of the catch-up of the foreign life expectancy theoretically.

It is important to note that our main finding is different from the impact of cross-country convergence in per capita income discussed in the literature (e.g., Chinn and Prasad, 2003; De Santis and Lührmann, 2009). The stages of development hypothesis for the balance of payments suggests that the catch-up of foreign income also has a quadratic polynomial relationship with the home current account balance, while this relationship is characterized by an "inverted" U-shaped curve. Specifically, a catching-up foreign country borrows against its future income and increases its current account deficit at the early stage of development, while this deficit decreases when reaching an advanced stage of development. As a result, a home country runs a bilateral current account surplus at first but eventually this bilateral balance deteriorates in the catching-up process of the foreign income. However, this is the opposite of the case of the U.S. current account balance with China, which is negative over the period after 2003 and has tended to improve in recent years as shown in Figure 2.[6] Therefore, our finding for the impact of life expectancy is helpful to interpret the fact that several developed countries including the United States run current account deficits against catching-up countries such as China.[7] Furthermore, we extend this discussion by suggesting that such deficits fall in the future if the life expectancy gap continues to narrow.

This paper is organized as follows. Section 2 briefly reviews the related literature. Section 3 explains the econometric methodology and the data. Section 4 presents our empirical results, and Section 5 checks their robustness. To provide theoretical support for the empirical results, Section 6 develops a two-country OLG model, and Section 7 presents the results of the comparative statics. Section 8 concludes this paper.

[5] In the United States, the labor force participation rate of those aged 65 years and over increased from 12.9% to 19.6% for the period 2000–2018. This growth is much faster than that in the 1990s, during which the participation rate increased from 11.8% to 12.3% (source: U.S. Bureau of Labor Statistics).

[6] Chinn and Prasad (2003) also point out that evidence to support the stages of development hypothesis is weak for 18 industrial and 71 developing countries.

[7] Several studies use the other factors to explain this fact. For example, Mendoza et al. (2009) focus on international financial integration and heterogeneous domestic financial markets. Gourinchas and Jeanne (2013) extend this discussion to capital flows across developing countries.

2. Brief review of the related literature

Several theoretical studies focus on life expectancy to examine demographic impacts on current account balances. Using the calibrated OLG model, Ferrero (2010) finds that a nonnegligible and nearly permanent component of the U.S. deficit with the other G7 countries can be described by demographic variables, especially the life expectancy gap. Similarly, Backus et al. (2014) and Ito and Tabata (2010) indicate the important role of life expectancy as a predictor of current account dynamics. However, in the empirical literature, little attention has been paid to the impact of the life expectancy gap. This paper contributes to the literature in this respect.

Since the seminal work of Chinn and Prasad (2003), many empirical studies have been conducted on the medium-term determinants of current account balances. In this literature, the old-age dependency ratio is regarded as a key determinant of national savings. Using global data sets, Chinn and Ito (2007), De Santis and Lührmann (2009), Gruber and Kamin (2007), and Lane and Milesi-Ferretti (2012) show that current account balances deteriorate in countries with a higher old-age dependency ratio. This paper extends the important contributions of these studies by extracting the impact of life expectancy from the old-age dependency ratio.

De Santis and Lührmann (2009) examine the determinants of net flows in equity securities and net flows in debt instruments (i.e., components of current account balances). They find that the old-age dependency ratio is associated with net inflows in equity securities and net outflows in debt instruments. Our approach is similar in terms of using disaggregated data. Although commonly used data on current account balances with all trading partners are useful for describing the direction and size of capital flows comprehensively, our empirical analysis based on bilateral data provides more specific information on the impact of Chinese life expectancy on the U.S. economy.

Time series and panel cointegration techniques are applied when current account balances and their determinants are found to be nonstationary. Using the time series data on the U.S. current account balance, Hung and Gamber (2010) estimate the cointegrating regression and error correction models based on the absorption approach. They find that the general dependency ratio (the sum of the youth and old-age dependency ratios) does not have a significant impact on the U.S. current account balance. Unger (2017) uses a panel of 11 member countries of the euro area. The cointegrating regression model for the current account balance includes credit growth as a key variable, and it is found that the long-run coefficient of the general dependency ratio is significantly negative. A similar econometric approach is used by Belke and Dreger (2013), who report that international competitiveness measured by real exchange rates is a key factor for explaining the regional imbalance in the euro area.

Arghyrou and Chortareas (2008) investigate nonlinear adjustment of current account balances toward long-run equilibrium relationships. They use the logistic smooth threshold error-correction model based on time series data and demonstrate that nonlinearity holds for the majority of euro area countries. As Choi and Saikkonen (2010) point out, nonlinear cointegration techniques are classified into two approaches: the first focuses on nonlinear adjustment mechanisms to deviations from long-run equilibrium relationships (e.g., Arghyrou and Chortareas, 2008); the second assumes that long-run equilibrium relationships themselves are nonlinear. Our analysis is based on the second approach.

3. Econometric methodology and data

3.1. Cointegrating polynomial regression

Our empirical analysis aims to demonstrate that the U-shaped relationship between the U.S. current account balance with China and the life expectancy gap in Figure 2 is a statistically valid result. For this, we use the polynomial cointegration technique developed by Wagner (2015) and Wagner and Hong (2016). This technique enables us to examine *nonlinear* long-run equilibrium relationships involving integrated regressors and their powers, and is particularly useful when economic theories predict that underlying variables have nonlinear relationships and/or linear models have poor predictive accuracy.

We estimate the polynomial regression model based on time series data as

$$z_t = \beta_0 + \beta_{11} x_{1t} + \beta_{12} x_{1t}^2 + \beta_2 x_{2t} + \cdots + \beta_m x_{mt} + u_t \,, \tag{1}$$

where $\beta_0, \cdots, \beta_m$ are coefficients to be estimated.[8] An explained variable z_t and explanatory variables in linear form $x_{1t}, \cdots, x_{mt}$ are integrated of order one. Only x_{1t} is allowed to have a polynomial relationship with z_t, and this is the empirically most relevant case as Wagner (2013) points out. Eq. (1) is regarded as a cointegrating polynomial regression if an error term

[8] The estimation results remain essentially unchanged if a time trend is included in the model.

u_t is stationary.

The existence of polynomial cointegration is a prerequisite for the successful modeling of the U-shaped relationship between the U.S. current account balance with China (z_t) and the life expectancy gap (x_{1t}) because the rejection of the stationarity of u_t means that Eq. (1) is a spurious regression. Note that standard tests for linear cointegration cannot be applied to Eq. (1) because powers of integrated variables are not integrated variables and thus x_{1t} and x_{1t}^2 cannot be separately treated as integrated regressors (Wagner, 2012). This problem is resolved by using the polynomial cointegration technique.

The testing procedure for polynomial cointegration involves two steps. First, the fully modified ordinary least squares (FMOLS) method developed by Phillips and Hansen (1990) is used to estimate Eq. (1), and the FMOLS residual $\hat{u}_t^+$ is obtained.[9] This estimation method is designed to correct for endogeneity and serial correlation biases nonparametrically. Next, we calculate the test statistic for the null hypothesis that u_t is stationary as

$$CT = \frac{1}{T\widehat{\omega}} \sum_{t=1}^{T} \left(\frac{1}{\sqrt{T}} \sum_{j=1}^{t} \hat{u}_j^+ \right)^2 , \qquad (2)$$

where T is the sample size and $\widehat{\omega}$ is a consistent estimator of the long-run variance of $\hat{u}_t^+$. Wagner and Hong (2016) show that the CT test statistic has a nuisance parameter-free limiting distribution and thus the corresponding critical values can be calculated in the special case where only one integrated regressor has a polynomial relationship with an explained variable, like x_{1t} in Eq. (1). The critical values are tabulated in Wagner (2013). If the CT test statistic is not significant, Eq. (1) is regarded as a long-run equilibrium relationship with a nonlinear impact of x_{1t}.

3.2. Data

The sample period is from the first quarter of 2003 to the fourth quarter of 2016 because the bilateral U.S. current account balance with China is available from 2003 and the

[9] Although Wagner and Hong (2016) propose the new FMOLS estimator for cointegrating polynomial regressions, Stypka et al. (2017) show that this estimator has the same asymptotic distribution as the standard FMOLS estimator. Therefore, both estimators are applicable to the test for polynomial cointegration. This paper uses the standard FMOLS method because it is easier to implement.

data on U.S. and Chinese life expectancy at birth are available until 2016 in the present study. Details of the data are reported in Appendix A. The analysis incorporates the U.S. current account balance, the life expectancy gap, and six control variables.

Current account balance

The bilateral U.S. current account balance with China is used as the explained variable. Following the existing literature, this variable is expressed as a percentage of U.S. GDP.

Life expectancy gap

The life expectancy gap is defined as U.S. life expectancy at birth minus Chinese life expectancy at birth. This variable is assumed to have a quadratic polynomial relationship to the U.S. current account balance with China. To increase the sample size, the annual data on the U.S. and Chinese life expectancy are converted into quarterly data by using the interpolation method provided by EViews. This method assigns each annual observation to the fourth quarter of the same year and then places all intermediate points on a natural cubic spline connecting all the points. This conversion seems to be a good approximation in the case of this study because, as shown in Figure 1, it is clear that the life expectancy for each country exhibits a strong trend and is less volatile around the trend. Indeed, the basic estimation results remain unchanged if we use the annual data (Section 4.2, Table 3).

Real domestic demand

Real domestic demand is an important determinant of the U.S. current account balance (e.g., IMF, 2018) and serves as a measure of business cycles. Nominal domestic demand is divided by the GDP deflator (the United States) or the consumer price index (China) to obtain real domestic demand. The data are converted into percentage changes from a year ago. Similar to the life expectancy gap, the Chinese data are subtracted from the U.S. data. This variable is expected to have a negative coefficient because weak domestic demand depresses imports and improves the current account balance (e.g., IMF, 2014; Lane and Milesi-Ferretti, 2012).

Real exchange rates

To measure international competitiveness, we use the real exchange rate. The nominal

USD/CNY exchange rate (units: Chinese yuan to one U.S. dollar) and the consumer price indices for the United States and China are used to calculate the real USD/CNY exchange rate. An increase in this variable means a real appreciation of the U.S. dollar and thus is negatively associated with the U.S. current account balance (e.g., Hung and Gamber, 2010). This variable is converted into natural logarithms.

Volatility index

For the first time in the literature, our cointegrating regression model includes the S&P 500 volatility index (VIX) to control for the effect of the global financial crisis. This variable is commonly used as a leading indicator of market risk. Its highest value was observed during the crisis of 2008. The inclusion of this variable is important because the financial crisis and the resulting increase in the market risk have a large impact on countries with large external deficits including the United States (Forbes and Warnock, 2012; Lane and Milesi-Ferretti, 2012). Given that a higher degree of market risk increases the risk aversion of investors, it reduces capital flows and improves the current account balance. Therefore, the expected sign of the coefficient on VIX is positive. The data are converted into natural logarithms.

Financial development

This paper uses stock market capitalization (SMC) as a ratio of GDP to measure financial development.[10] The U.S. data are divided by the Chinese data and then converted into natural logarithms. Financial development involves reducing information acquisition and transaction costs, overcoming or managing information asymmetries, and improving corporate governance (Ito and Chinn, 2009). Hence, it promotes both savings and investment, and its coefficient is regarded as a net impact (Chinn and Prasad, 2003). For example, Gruber and Kamin (2009) show that the SMC/GDP ratio is positively associated with current account balances in industrial countries.

Working-age population

[10] If we use the M2/GDP ratio instead of the SMC/GDP ratio, the estimation results for the impact of the life expectancy gap remain unchanged. However, we also find that this adjustment deteriorates the model's performance. The results are not reported in this paper but are available from the author upon request.

Working-age population (aged 15–64 years) is included in our regression model as another demographic factor.[11] The U.S. data are available in quarterly frequency, and the annual Chinese data are converted into quarterly data by using the same interpolation method as explained above. The quarterly data are used to calculate percentage changes from a year ago, and then the Chinese data are subtracted from the U.S. data. Gruber and Kamin (2007) suggest that the working-age population obtains wage income and increases its savings, and therefore, the cohort size has a positive coefficient.

Fiscal balance

To control for the impact of the government sector, we use the fiscal balance as a ratio of GDP. The U.S. data are available in quarterly frequency, and the annual Chinese data are converted into quarterly data by using the same interpolation method as explained above and then are subtracted from the U.S. data.[12] This variable is expected to have a positive coefficient because, in the absence of a full Ricardian offset, an increase in the fiscal balance leads to an increase in national savings (Chinn and Prasad, 2003).

4. Empirical results

4.1. Unit root tests

Before estimating Eq. (1), it is necessary to examine the stationarity of each of the variables introduced in Section 3 because the polynomial cointegration technique is applicable when the variables are integrated of order one. This paper uses the unit root test developed by Elliott et al. (1996), which is a modification of the Dickey and Fuller (1979) test.

Table 1 around here

The results are reported in Table 1. The test does not reject the null hypothesis of a unit root for the variables in levels. However, the null hypothesis of a unit root is rejected for the variables in first differences. Therefore, all the variables are integrated of order one. This

[11] If we use the fertility rate instead of the growth rate of the working-age population, the main results remain unchanged. The results are not reported in this paper but are available from the author upon request.

[12] Unlike in the cases of the demographic variables, this conversion may not be a good approximation because the fiscal balance/GDP ratio for China is relatively volatile. However, if the Chinese data are not subtracted from the U.S. data as in Hung and Gamber (2010), the main results remain unchanged. The results are not reported in this paper but are available from the author upon request.

finding is consistent with previous studies (e.g., Arghyrou and Chortareas, 2008).

4.2. Polynomial cointegration

We can now examine the presence of cointegration for Eq. (1), in which the life expectancy gap (x_{1t}) has a quadratic polynomial relationship to the U.S. current account balance with China (z_t). Following Wagner (2013, 2015), the CT test statistic for the null hypothesis of polynomial cointegration is calculated from the basic model for $m = 1$.[13]

The test result is shown in Table 2. The CT test statistic is well below the 5% critical value of 0.2927, showing that polynomial cointegration for Eq. (1) holds. Therefore, a nonlinear long-run equilibrium relationship exists between the U.S. current account balance with China and the life expectancy gap.[14]

Table 2 around here

As explained in Section 3.1, the FMOLS method is used to estimate the nonlinear long-run equilibrium relationship. Following Wagner (2015), we also use the dynamic ordinary least squares (DOLS) method developed by Saikkonen (1991) and Stock and Watson (1993) for reference purposes.[15] The estimation results for the basic and full models are reported in Table 3. These results do not suffer from endogeneity and serial correlation biases because the FMOLS and DOLS methods are designed to correct for these biases. The coefficients on the life expectancy gap and its square are significantly negative and positive, respectively. The estimation results are robust if we use the annual data.[16] Therefore, the long-run equilibrium

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[13] Although our full model consists of 7 integrated regressors including the life expectancy gap ($m = 7$) as explained in Section 3, the critical values for the models with $m > 4$ are not reported in Wagner (2013). However, as shown below, the estimation results of the coefficients on the life expectancy gap and its square are little affected even though we add the other 6 integrated regressors in the basic model.

[14] We also find that the growth rate difference in working-age population does not have a nonlinear impact on the U.S. current account balance with China (Table 2a). Therefore, although the old-age dependency ratio is decomposed into life expectancy and the growth rate of working-age population, only the former has a nonlinear impact. The latter has a significant linear impact (Table 3 below).

[15] The number of leads and lags in the DOLS regression is 1. For the extension of the DOLS method into nonlinear cointegration, see Choi and Saikkonen (2010).

[16] The full model is not estimated when using the annual data because the sample size is very small. The DOLS regression based on the annual data excludes the lead term because it is preferable that the end of the effective sample period is 2016 for consistency with the other results. If the lead term is included, the coefficients on the life expectancy gap and its square are significant and the threshold level is 2.73.

relationship between the U.S. current account balance with China and the life expectancy gap is described by the U-shaped curve. The fitted line of the basic model is shown in Figure 2. Robustness checks are conducted in Section 5.

Table 3 around here

The threshold level of the life expectancy gap in Table 3 is calculated as $-\beta_{11}/2\beta_{12}$ and is approximately 3. Hence, the sign of the impact of the life expectancy gap is reversed when the gap narrows by three years or less, and this case is applied to the period after 2013. In other words, the catch-up of Chinese life expectancy worsens the U.S. deficit until 2012, but thereafter the increased U.S. deficit is improved by a further catch-up of Chinese life expectancy. More detailed discussion on the threshold level of the life expectancy gap is provided in Section 5.2.

These findings are obtained after controlling for major determinants of current account balances. The estimation results for the full model show that the U.S. current account balance is improved by weak domestic demand, a real depreciation of the U.S. dollar, a higher market risk, a higher degree of financial market development, increased working-age population, and an improved fiscal balance. These results are in line with the empirical literature (e.g., Belke and Dreger, 2013; Chinn and Prasad, 2003; De Santis and Lührmann, 2009; Gruber and Kamin, 2007, 2009; Lane and Milesi-Ferretti, 2012; Unger, 2017).

4.3. Forecast performance

We find that the U-shaped impact of the life expectancy gap is a statistically valid result because polynomial cointegration holds and the coefficient on the squared term of the life expectancy gap is significant after controlling for the major determinants of current account balances and after correcting for endogeneity and serial correlation biases. To further examine the importance of allowing for the U-shaped impact, we compare the forecast performance of the linear and nonlinear models. Specifically, by imposing the restriction that $\beta_{12} = 0$ on Eq. (1), a standard linear regression model is obtained as

$$z_t = \beta_0 + \beta_{11}x_{1t} + \beta_2 x_{2t} + \cdots + \beta_m x_{mt} + u_t\,. \tag{3}$$

A performance comparison of Eqs. (1) and (3) enables us to understand how much the nonlinear impact of the life expectancy gap contributes to the improvement of the forecast accuracy.

In the forecast analysis, the estimation period is from the first quarter of 2003 to the fourth quarter of 2015, and the FMOLS method is used to estimate the models with the linear and nonlinear specifications of the impact of the life expectancy gap.[17] The forecast period is from the first quarter of 2016 to the fourth quarter of 2016, and the estimated coefficients and the data on the explanatory variables are used to calculate the out-of-sample forecast values of the U.S. current account balance. The evaluation of forecast accuracy is based on the root mean squared error (RMSE).[18]

The results are reported in Table 4. The values of the RMSE are 0.601 and 0.157 for the linear and nonlinear full models, respectively, indicating that the RMSE for the full model is decreased by 74% just by adding the squared term of the life expectancy gap. A similar observation applies to the basic model. Therefore, the forecast accuracy for the U.S. current account balance with China is dramatically improved by allowing for nonlinearity in the impact of the life expectancy gap.

Table 4 around here

The poor predictive performance of the linear models is possibly due to omitted variable problems. For example, we find that the U.S. current account balance with China and the life expectancy gap are not cointegrated in the linear basic model.[19] Therefore, the linear specification is not suitable for describing their long-run equilibrium relationship. Furthermore, for the linear full model, the control variables, except for the fiscal balance/GDP ratio and VIX, are not significant, suggesting the possibility of model misspecification.[20] These problems are resolved by adding the squared term of the life expectancy gap in the models.

5. Discussion

5.1. Robustness checks

To check the robustness of the empirical results, we estimate Eq. (1) with all possible combinations of the control variables, and by doing so, we can obtain 64 estimation results for

[17] If we use the data for this subsample period, the estimation results for Eq. (1) reported in Table 3 remain essentially unchanged (see Table 4a).

[18] The results for the forecast evaluation remain unchanged if we use alternative criteria such as the mean absolute percentage error and the Theil inequality coefficient (see Table 4b).

[19] The Engle and Granger (1987) approach is used as a test for linear cointegration. The computed test statistic is –2.247, suggesting the absence of linear cointegration.

[20] The results are not reported in this paper but are available from the author upon request.

the nonlinear impact of the life expectancy gap.[21] The data are the same as in Section 4.2.

The results of the FMOLS estimation are summarized in Table 5.[22] For all cases, the coefficients on the life expectancy gap and its square are significantly negative and positive, respectively, and the threshold level of the life expectancy gap is approximately 3. These results are consistent with those reported in Table 3. Therefore, there is sufficient evidence in favor of the U-shaped impact of the life expectancy gap.

Table 5 around here

5.2. Savings rate in China

As Figure 1 shows, the life expectancy gap has fallen over time and has been less than the threshold level of approximately 3 since 2013. Hence, we find that the catch-up of Chinese life expectancy changes the direction of its impact from 2013. To interpret this result, we focus on the savings rate (net household savings as a percentage of household disposable income) in China. Given that Chinese life expectancy has a nonlinear impact on capital export to the United States, it should also have a similar nonlinear impact on savings.

The data on the savings rate and life expectancy at birth in China are used. The sample period is from 2003 to 2016 to maintain consistency with the empirical analysis in Section 4. Because these data are available only in annual frequency and the sample size is very small, the results of this analysis are only suggestive. The polynomial cointegration technique is used again, and the savings rate is regressed on life expectancy, its square, and a constant term. The CT test statistic for the null hypothesis of quadratic polynomial cointegration is 0.227 and is not significant at the 5% level.

The estimation results are reported in Table 6. We find that the relationship between the savings rate and life expectancy in China is characterized by an *inverted* U-shaped curve. The fitted line is shown in Figure 3. Therefore, given the continued increase in Chinese life expectancy over the sample period, the savings rate increases at first and then decreases.

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[21] Because six control variables are used, the total number of combinations is counted as follows: (a) $6C0 = 1$, (b) $6C1 = 6$, (c) $6C2 = 15$, (d) $6C3 = 20$, (e) $6C4 = 15$, (f) $6C5 = 6$, and (g) $6C6 = 1$. The estimation results for (a) and (g) are reported in Table 3.

[22] The estimation results for the coefficients on the control variables are not reported in Table 5 for space reasons but are available from the author upon request. The main results remain essentially unchanged if we use the DOLS method (see Table 5a).

Furthermore, Chinese life expectancy has been more than the threshold level of 75.63 since 2013. These results are consistent with our main findings for the U-shaped impact of the life expectancy gap. Specifically, the catch-up of Chinese life expectancy increases the domestic savings rate, and this effect promotes capital export to the United States. However, the savings rate starts to decrease as Chinese life expectancy exceeds the threshold level. Consequently, the narrowing life expectancy gap now reduces capital export to the United States.

Table 6 around here

Figure 3 around here

5.3. Estimation of error correction models

Several empirical studies examine adjustment mechanisms to deviations from long-run equilibrium relationships (Arghyrou and Chortareas, 2008; Hung and Gamber, 2010; Unger, 2017). Along this line, we estimate the error correction model as

$$\Delta z_t = \mu + \psi EC_{t-1} + \pi_0 \Delta z_{t-1} + \pi_1 \Delta x_{1,t-1} + \cdots + \pi_m \Delta x_{m,t-1} + e_t \, , \qquad (4)$$

where Δ denotes the difference operator; μ, a constant term; EC_t, an error correction term; ψ, an adjustment coefficient; $\pi_0, \ldots, \pi_m$, short-run coefficients; and e_t, an error term. The error correction term is calculated as the FMOLS residual in Eq. (1), and the coefficients reported in Table 3 are used. Hence, $EC_t > 0$ means that z_t (the U.S. current account balance with China in period t) is above its long-run equilibrium value, and vice versa. Although the lag length of the error correction model is assumed to be 1, the estimation results for the adjustment mechanism are robust if the lag length is set to 2 and 3.[23] The data are the same as in Section 4.2.

For consistency with the basic and full models used in the cointegration analysis, the number of integrated regressors m is set to 1 and 7. The estimation results for Eq. (4) are reported in Table 7. For both specifications, the assumptions of residual normality and no residual autocorrelation hold, suggesting that Eq. (4) is successfully estimated. The adjustment coefficient is significantly negative, and the estimated values of –0.244 and –0.311 indicate that the error correction mechanism works appropriately. Therefore, the dynamics of the U.S.

[23] The diagnostic test results reported below suggest that the model with one lag is adequate for describing the dynamics of the U.S. current account balance with China.

current account balance with China are affected by the deviations from the nonlinear long-run equilibrium relationship. Although the autoregressive coefficient π_0 is significant, the other short-run coefficients $(\pi_1, \ldots, \pi_7)$ are not significant except for the SMC/GDP ratio, suggesting that the long-run impact is dominant.[24]

Table 7 around here

6. Theoretical model

The evidence for the U-shaped impact of the life expectancy gap is new in the literature, and thus a theoretical foundation for the mechanism of the impact is required. For this purpose, we develop a simple and analytically tractable OLG model. Our model is based on Ito and Tabata (2010), who show an analytical result for the impact of life expectancy on current account balances. To extend their model, this paper uses model elements developed by Aísa et al. (2012) and Ludwig and Vogel (2010) and incorporates elderly labor supply into the model.

Krueger and Ludwig (2007) also use an open-economy OLG model with elderly labor supply, and present simulation results for the welfare effects of extended retirement ages. Although our model is in line with their model, a welfare analysis is not conducted in this paper because our focus is on the U-shaped impact of the life expectancy gap. Instead, for the first time in this literature, we derive an analytical expression for the general equilibrium effect of foreign life expectancy on the home current account balance consisting of three components, namely savings for retirement, taxes, and elderly productivity.

6.1. Basic setup

We use a two-country, one-good, two-period OLG model with lifetime uncertainty. The world consists of a home country and a foreign country, and they are linked through an integrated commodity market and an international capital market. However, the domestic labor market is closed. Although the structure of the home country is explained in this section, the structure of the foreign country can be discussed in the same way as that of the home country. Foreign variables and parameters are denoted by an asterisk.

[24] For the model with $m = 7$, the Wald test statistic for the null hypothesis that $\pi_1 = \cdots = \pi_7 = 0$ is 11.544 and its corresponding p-value is 0.117.

Our model assumes that a new generation, called generation t, is born in each period $t = 1, 2, 3, \dots$. Generation t is composed of a continuum of $N_t > 0$ units of identical agents. In period 1, there are N_0 units of initial old agents who live only in period 1. Let $n > 0$ be the population growth rate, and thus $N_t = (1 + n)N_{t-1}$. The lifetime of agents is uncertain and they live for a maximum of two periods, young and old age. An agent in the home country survives to old age with a probability of $p \in [0, 1]$, and dies at the beginning of old age with a probability of $1 - p$. Therefore, in each period, there are N_t units of young agents and pN_{t-1} units of old agents. A rise of the survival probability p leads to population aging in the home country because it increases the share of old agents. Hence, the survival probability corresponds to life expectancy (e.g., Aísa et al., 2012).

Under these assumptions, the old-age dependency ratio is given by

$$\frac{pN_{t-1}}{N_t} = \frac{p}{1+n}. \tag{5}$$

Therefore, the effect of the old-age dependency ratio can be decomposed into the effects of life expectancy and the growth rate of working-age population.[25] For this reason, our empirical analysis uses these two variables instead of the old-age dependency ratio.

To describe the catching-up process of foreign life expectancy theoretically, we restrict our analysis to the case where the parameter condition $p^* < p$ holds. We also assume that the home and foreign countries have the same population growth rate. This assumption is helpful in obtaining clear analytical results for the impact of the life expectancy gap.[26]

6.2. Households

Households derive utility from their own consumption in young and old age. In this paper, the expected lifetime utility of the representative agent in generation t is defined as

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[25] As mentioned below, we assume that all young agents work. Hence, the population growth rate is identical to the growth rate of working-age population in this model. The main empirical results remain unchanged if we use the fertility rate instead of the growth rate of working-age population (see footnote 11).

[26] As Ito and Tabata (2010) point out, if the population growth rate differs between the two countries, the low-fertility country becomes a small country in the steady-state equilibrium and its demographic changes are not reflected in world factor prices. This implies that life expectancy in the low-fertility country does not affect international capital flows, and our theoretical analysis becomes complicated. For the period 2003–2016, the average annual growth rates of the working-age population for the United States and China are 0.79% and 0.76%, respectively.

$$u_t = \ln(c_{1,t}) + p\ln(c_{2,t+1}) , \tag{6}$$

where $c_{1,t}$ and $c_{2,t+1}$ denote consumption in young and old age, respectively.

Young agents in generation t have one unit of time endowment. They inelastically supply one unit of effective labor and obtain wage income w_t. A part of the wage income is devoted to current consumption $c_{1,t}$ and the payment of taxes $\tau_t w_t$, where τ_t is the rate of tax levied on wage income, and the rest is savings s_t.

Following Ito and Tabata (2010), we assume the existence of an actuarially fair insurance company developed by Yaari (1965). This company collects funds and invests them in firms or foreign countries. Because individuals do not migrate, the insurance contracts differ between the two countries. Returns on investment are distributed uniformly among insured surviving old agents. Therefore, those in the home and foreign countries receive payments $\frac{R_{t+1}}{p} s_t$ and $\frac{R_{t+1}}{p^*} s_t^*$ in exchange for s_t and s_t^*, respectively.[27] R_{t+1} is the gross rate of interest, and the insurance contract is more attractive to agents as long as the survival probability is less than 1. Consequently, in the absence of a bequest motive, agents prefer to invest their assets in the insurance.

Old agents in generation t also have one unit of time endowment and consume their entire wealth. Following Ludwig and Vogel (2010), this paper assumes that old agents work a given fraction ω of their time. For the rest of the time $(1-\omega)$, they are retired and receive pension benefits b_{t+1}. Hence, ω can be regarded as a measure of retirement age. Another interpretation of ω is that it is the proportion of old agents who work (i.e., the labor force participation rate of the elderly) because of the assumption of identical agents. Either way, ω works like a policy variable, and a change in ω is interpreted as a change in labor market incentives for the elderly (Ludwig and Vogel, 2010).

Let δ be the fraction of young productivity maintained in the second period of life. Old agents in generation t inelastically supply $\omega\delta$ units of effective labor, obtain wage income $\omega\delta w_{t+1}$, and pay taxes $\tau_{t+1}\omega\delta w_{t+1}$. In this paper, δ is specified as

$$\delta = p^\gamma , \tag{7}$$

[27] The coexistence of the actuarially fair private annuity and the publicly provided social security can be justified as a hedging strategy against different risks in the capital and labor markets that are not modeled here explicitly (Ito and Tabata, 2010).

where γ is a positive parameter. This specification is essentially the same as that developed by Aísa et al. (2012) and Hirazawa and Yakita (2017). Because $\delta \leq 1$, productivity in old age is lower than in the young. However, the productivity decline in old age is mitigated by an increase in life expectancy p (i.e., the improvement of health). This phenomenon is consistent with the theory of the compression of morbidity developed in the field of health and medical sciences (e.g., Fries, 1980), and we show that this assumption holds empirically (Appendix B).

Based on these settings, the lifetime budget constraints of the representative agent in generation t are expressed as

$$c_{1,t} + s_t = (1 - \tau_t) w_t \,, \tag{8}$$

$$c_{2,t+1} = \frac{R_{t+1}}{p} s_t + (1 - \tau_{t+1}) \omega \delta w_{t+1} + (1 - \omega) b_{t+1} \,. \tag{9}$$

Maximizing Eq. (6) subject to Eqs. (8) and (9) gives the saving equation as

$$s_t = \frac{p}{1+p} \left[(1 - \tau_t) w_t - \left\{ \frac{(1 - \tau_{t+1}) \omega \delta w_{t+1} + (1 - \omega) b_{t+1}}{R_{t+1}} \right\} \right] . \tag{10}$$

A higher life expectancy p increases savings because it means a higher probability of enjoying consumption in old age. However, a higher earning ability in old age δ decreases savings, which suggests that young agents who expect to earn more in old age do not have incentives to save more.

6.3. Government

The government runs a social security system. The government budget is assumed to be balanced in each period. Therefore, total contributions by workers are equal to total pension payments as

$$\tau_t w_t N_t + \tau_t w_t \omega \delta p N_{t-1} = b_t (1 - \omega) p N_{t-1} \,. \tag{11}$$

Following Bárány et al. (2018), Ito and Tabata (2010), Krueger and Ludwig (2007), and Ludwig and Vogel (2010), our model incorporates a pension scheme where the payments are calculated using a replacement rate. Let us assume that $b_t = \phi(1 - \tau_t) w_t$, where $\phi \in [0,1)$ is the replacement rate.[28] Then, the tax rate is given by

[28] Ito and Tabata (2010) assume that $b_t = \phi w_t$. If we use this assumption, the U-shaped relationship between the home current account balance and the life expectancy gap can be derived.

$$\tau_t = \frac{(1-\omega)\phi p}{1+n+\omega\delta p+(1-\omega)\phi p} \equiv \tau\,. \tag{12}$$

From this equation, we have

$$\frac{d\tau}{dp} = \underbrace{\frac{\partial\tau}{\partial p}}_{+} + \underbrace{\overbrace{\frac{\partial\tau}{\partial\delta}}^{-}\overbrace{\frac{\partial\delta}{\partial p}}^{+}}_{-}\,. \tag{13}$$

An increase in life expectancy raises the tax rate because it increases the number of pensioners, while an improvement of elderly productivity eases the tax burden per worker and decreases the tax rate. Using Eqs. (7) and (12), we obtain the following.

$$\frac{d\tau}{dp} = \frac{(1+n-\gamma\omega\delta p)(1-\omega)\phi}{[1+n+\omega\delta p+(1-\omega)\phi p]^2} > 0 \tag{14}$$

Life expectancy is positively associated with the tax rate unless γ in Eq. (7) is large enough. The numerical analysis in Section 7 shows that the positive sign in Eq. (14) holds.[29]

6.4. Firms

Firms produce output using an aggregate technology that takes a standard Cobb–Douglas form as

$$Y_t = \bar{A}K_t^{\alpha}L_t^{1-\alpha}\,, \tag{15}$$

where Y_t denotes aggregate output; $\bar{A}$, a technology parameter; K_t, aggregate capital; L_t, aggregate labor; and $\alpha \in (0,1)$. Following Ludwig and Vogel (2010), L_t is given by the sum of effective labor inputs supplied by young and old agents as

$$L_t = N_t + \omega\delta p N_{t-1}\,. \tag{16}$$

Note that $L_t = (1+n)L_{t-1}$ because Eq. (16) can be rewritten as $L_t = \left(1+\frac{\omega\delta p}{1+n}\right)N_t$.

The existing capital stock is fully used for old-age consumption and the gross interest rate R_t contains a 100% depreciation rate. Assuming that firms operate in perfect competition, the first-order conditions for profit maximization are

$$w_t = (1-\alpha)\bar{A}k_t^{\alpha}\,, \tag{17}$$

$$R_t = \alpha\bar{A}k_t^{\alpha-1}\,, \tag{18}$$

[29] The estimated value of γ is 3.195 and is sufficiently small to satisfy the condition that $1+n > \gamma\omega\delta p$. Therefore, this paper assumes that this condition holds.

where $k_t \equiv \frac{K_t}{L_t}$.

6.5. Equilibrium

Assuming free mobility of capital, the rates of return on capital in the home and foreign countries converge to one world price R_t^W, where $R_t^W = R_t = R_t^*$. This means that $k_t = k_t^*$ and $w_t = w_t^*$ from Eqs. (17) and (18).

Let $N_t^W \equiv N_t + N_t^*$ and $L_t^W \equiv L_t + L_t^*$. Because N_t, N_t^*, L_t, and L_t^* grow at the same rate n, we have $N_t^W = (1+n)N_{t-1}^W$ and $L_t^W = (1+n)L_{t-1}^W$. The market clearing condition for world capital is $K_{t+1}^W = s_t N_t + s_t^* N_t^*$, where $K_{t+1}^W \equiv K_{t+1} + K_{t+1}^*$. Dividing both sides of this condition by N_t^W gives

$$k_{t+1}^W[\varepsilon\theta + \varepsilon^*(1-\theta)] = s_t\theta + s_t^*(1-\theta) \equiv s_t^W , \tag{19}$$

where $k_t^W \equiv \frac{K_t^W}{L_t^W}$, $\varepsilon \equiv 1 + n + \delta\omega p$, $\varepsilon^* \equiv 1 + n + \delta^*\omega^* p^*$, and $\theta \equiv \frac{N_t}{N_t^W} = \frac{N_0}{N_0^W}$. The population share of the home country θ is derived using the facts that $N_t = (1+n)^t N_0$ and $N_t^* = (1+n)^t N_0^*$. From Eq. (19), we have $k_t^W = \frac{k_t\varepsilon\theta + k_t^*\varepsilon^*(1-\theta)}{\varepsilon\theta + \varepsilon^*(1-\theta)}$. Therefore, $k_t = k_t^* = k_t^W$ holds when $k_t = k_t^*$.

Let us assume that firms in the home and foreign countries have the same production technology. Then, substituting (10), (12), (17), (18), $b_t = \phi(1-\tau_t)w_t$, and those of the foreign country counterparts into Eq. (19) and taking the fact that $k_t = k_t^* = k_t^W$ into account, we can describe the dynamics of k_t^W as

$$k_{t+1}^W = \Gamma(k_t^W)^\alpha , \tag{20}$$

where

$$\Gamma \equiv \frac{\theta\lambda + (1-\theta)\lambda^*}{\theta\mu + (1-\theta)\mu^*},$$

$$\lambda \equiv \frac{p}{1+p}(1-\tau)(1-\alpha)\bar{A} ,$$

$$\mu \equiv \frac{p}{1+p}\frac{1-\alpha}{\alpha}(1-\tau)[\omega\delta + (1-\omega)\phi] + \theta\varepsilon + (1-\theta)\varepsilon^* ,$$

$$\lambda^* \equiv \frac{p^*}{1+p^*}(1-\tau^*)(1-\alpha)\bar{A} ,$$

$$\mu^* \equiv \frac{p^*}{1+p^*}\frac{1-\alpha}{\alpha}(1-\tau^*)[\omega^*\delta^* + (1-\omega^*)\phi^*] + \theta\varepsilon + (1-\theta)\varepsilon^* .$$

Therefore, the steady-state capital–labor ratio in an open economy k^W is given by

$$k^W = \Gamma^{\frac{1}{1-\alpha}}. \tag{21}$$

Eq. (21) can be used to examine the long-run equilibrium in closed economies. Specifically, if $\theta = 1$, the world economy consists of only the agents of the home country. In this case, Eq. (21) is interpreted as the steady-state capital–labor ratio in the closed home economy k_c, which is expressed as

$$k_c = \Psi^{\frac{1}{1-\alpha}}, \tag{22}$$

where $\Psi \equiv \frac{\lambda}{\tilde{\mu}}$ and $\tilde{\mu} \equiv \frac{p}{1+p}\frac{1-\alpha}{\alpha}(1-\tau)[\omega\delta + (1-\omega)\phi] + \varepsilon$. Similarly, when $\theta = 0$, the steady-state capital–labor ratio in the closed foreign economy k_c^* is determined as

$$k_c^* = \Psi^{*\frac{1}{1-\alpha}}, \tag{23}$$

where $\Psi^* \equiv \frac{\lambda^*}{\tilde{\mu}^*}$ and $\tilde{\mu}^* \equiv \frac{p^*}{1+p^*}\frac{1-\alpha}{\alpha}(1-\tau^*)[\omega^*\delta^* + (1-\omega^*)\phi^*] + \varepsilon^*$.

6.6. Current account balances

When the economies are opened to trade, the current account balance in the home country equals the change in net foreign assets as

$$G_t = (A_{t+1} - K_{t+1}) - (A_t - K_t), \tag{24}$$

where $A_t \equiv s_{t-1}N_{t-1}$ denotes the wealth in the home country, which includes opportunities to acquire claims on foreign capital (Geide-Stevenson, 1998). Using Eqs. (8), (9), (11), and the assumption of perfect competition, Eq. (24) can be rewritten as

$$G_t = Z_t + (R_t - 1)(A_t - K_t), \tag{25}$$

where $Z_t \equiv Y_t - c_{1t}N_t - c_{2t}pN_{t-1} - K_{t+1}$ is the trade balance in the home country. Hence, the current account balance in this model consists of the trade balance and net primary income.

Let $g_t \equiv \frac{G_t}{L_t}$. Then, the steady-state value of the current account balance per unit of effective labor g is given by

$$g = n(a - k^W), \tag{26}$$

where $a \equiv \frac{s}{\varepsilon}$, $s \equiv s(k^W, k^W)$, and $s(k_t, k_{t+1}) = \lambda(k_t)^\alpha - (\tilde{\mu} - \varepsilon)k_{t+1}$.[30]

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[30] If we use $\frac{G}{Y}$ instead of g, the following results remain essentially unchanged. The results are not reported in this paper but are available from the author on request.

7. Comparative statics

7.1. Analytical results

We can now conduct a comparative statics analysis in the steady state. The pattern of international capital flows is determined by the difference in the capital–labor ratio in the closed economies. For example, $k_c^* > k_c$ means $R^* < R$ from Eq. (18). Thus, after the economies are opened to trade, the young agents in the foreign country invest part of their savings in the home country to obtain a higher rate of return. The home country imports capital from the foreign country and runs a current account deficit in its long-run equilibrium (i.e., $g < 0$). Similarly, $g > 0$ holds when $k_c^* < k_c$. Therefore, to examine the impact of p^* on g, we first calculate a change in k_c^* based on the derivative of k_c^* with respect to p^*. Then, the resulting level of k_c^* is compared with the level of k_c. This procedure is the same as used in Ito and Tabata (2010).

Taking the derivative of k_c^* with respect to p^* gives

$$\frac{dk_c^*}{dp^*} = \underbrace{\frac{\partial k_c^*}{\partial p^*}}_{+} + \underbrace{\overbrace{\frac{\partial k_c^*}{\partial \tau^*}}^{-}\overbrace{\frac{d\tau^*}{dp^*}}^{+}}_{-} + \underbrace{\overbrace{\frac{\partial k_c^*}{\partial \delta^*}}^{-}\overbrace{\frac{d\delta^*}{dp^*}}^{+}}_{-}. \tag{27}$$

The proof is provided in Appendix C. This equation shows that foreign life expectancy affects the home current account balance through three channels. From Eq. (10), a rise in the savings rate increases the capital–labor ratio (first term: saving effect). However, from Eq. (14), an increased tax rate leads to a lower capital–labor ratio (second term: tax effect). Similarly, from Eqs. (7) and (10), an improvement of elderly productivity decreases the capital–labor ratio (third term: productivity effect). Consequently, the total effect $\frac{dk_c^*}{dp^*}$ varies in accordance with the sizes of these three effects.

7.2. Numerical analysis based on analytical results

This paper adopts a numerical comparative static analysis because it is very difficult to specify the sign of the total effect in Eq. (27) analytically. The parameters are reported in Table 8 and are based on the data for the United States and China. Using these parameters, we calculate the first, second, and third derivatives in Eq. (27) numerically, and their analytical expressions are shown in Appendix C. To describe the catching-up process of foreign life expectancy, the initial conditions are given by $p = 0.85$ and $p^* = 0.80$ and then only p^* is increased up to 0.84.[31]

Table 8 around here

The results of the numerical comparative statics are presented in Table 9. The total effect is positive for $p^* \leq 0.82$ but negative for $p^* > 0.82$.[32] Hence, k_c^* increases at first but then decreases when the life expectancy gap $p - p^*$ is narrowed by an increase in p^*. Corresponding changes in the home current account balance (per unit of effective labor) g are described in Figure 4, in which $p - p^*$ has a U-shaped relationship with g. Using these results, we can provide a theoretical foundation for the impact of the life expectancy gap as follows.

Table 9 around here

Figure 4 around here

At first, the saving effect is dominant (Table 9) and thus k_c^* increases as p^* increases, which means that an increase in p^* leads to an increase in s^* from Eq. (10) and this promotes capital accumulation. A subsequent decrease in R^* leads to larger capital flows to the home country, and thus g decreases as p^* starts to catch up with p (point A in Figure 4). Next, g reaches the turning point when p^* is approximately 0.825 (point B in Figure 4). Finally, k_c^* decreases in the middle of the convergence process of p^* because the sum of the tax and productivity effects exceeds the saving effect as shown in Table 9. Consequently, a further narrowing of the gap $p - p^*$ improves g (point C in Figure 4).

An important fact is that the male survival rate to age 65 years for China exceeded the threshold level of 0.825 in 2013. This result is consistent with the empirical results showing that the life expectancy gap reached its threshold level in 2013 (Section 4.2), Therefore, the U-shaped impact of the life expectancy gap can be theoretically reproduced.

7.3. Role of elderly labor supply

As mentioned in Section 6, we extend the theoretical literature by incorporating the

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[31] The male survival rate to age 65 years for China was 79.46% in 2003 and increased to 83.54% in 2016 (source: World Development Indicators of the World Bank).

[32] This result remains unchanged if we calculate the total effect for the parameter region $0.10 \leq p^* \leq 1.00$. Hence, the total effect is positive for $0.10 \leq p^* \leq 0.82$ and is negative for $0.82 < p^* \leq 1.00$.

productivity effect in Eq. (27) into the model. This extension is useful for describing the reversal of g because the comparative static analysis in Table 9 shows that the saving effect is larger in absolute value than the tax effect. To obtain clearer results for this point, g is calculated under the assumption that $\omega = \omega^* = 0$. This means that elderly labor supply and their productivity (and thus, the productivity effect) are excluded from the model, and p^* affects g through the remaining two channels. The other parameters are the same as in Table 8.

The calculation results for g with the assumption that $\omega = \omega^* = 0$ are presented in Figure 5. We find that g is not reversed even though the gap $p - p^*$ is narrowed by an increase in p^*, and this finding is inconsistent with the empirical evidence. Therefore, in the framework of our OLG model, the productivity effect is essential for describing the U-shaped impact of the life expectancy gap.

Figure 5 around here

Although the direct impacts of ω and ω^* on g are also of interest, the empirical analysis for those impacts is very difficult because the data on the labor force participation rate of the elderly for China are available per 10-year period and its sample size is 2.[33] From a theoretical perspective, our model suggests that individuals who plan to work longer in old age save less over their working life because they anticipate a shorter retirement, and an increase in elderly labor supply can therefore be associated with an increase in capital import and consequent deterioration of current account balances. The labor force participation rate of the elderly was 17.4% for the United States in 2016 and 21.1% for China in 2010, and the simulation results for the impact of $p - p^*$ on g remain almost unchanged if we assume that $\omega = 0.174$ and $\omega^* = 0.211$.[34]

7.4. Transitional analysis

In addition to the steady-state analysis conducted above, we examine the dynamic transition path of the home current account balance. The procedure of this analysis involves two steps. First, we assume that the world economy is in a steady-state equilibrium where $p =$

[33] The data are available only in 2000 and 2010 (source: OECD).

[34] These results are not reported in this paper but are available from the author upon request.

0.85 and $p^* = 0.80$ for $t < 11$. Next, p^* is increased to 0.84 for $t \geq 11$ to maintain consistency with the steady-state analysis, and the transition path of g_t toward the new steady-state equilibrium is described. The other parameters are the same as in Table 8.

The simulation results are indicated in Figure 6. We find that g_t deteriorates at first but eventually improves over time as the life expectancy gap narrows. Hence, the transition dynamics of g_t are consistent with the empirical results. Specifically, the increase in p^* motivates foreign young agents in period 11 to save more for their retirement. However, their social security burden does not increase because the foreign life expectancy in generation 10 is unchanged. Therefore, in period 12, the foreign capital–labor ratio increases, and the home current account balance decreases. For the subsequent periods, the social security burden and effective labor input supplied by old agents increase, and the aggregate effect of taxes and elderly productivity leads to a decrease in the foreign capital–labor ratio and thus an increase in the home current account balance.

Figure 6 around here

8. Conclusions

There is sufficient evidence in favor of the U-shaped impact of the life expectancy gap on the U.S. current account balance with China, and the threshold level of the life expectancy gap is successfully estimated. Therefore, the catch-up of Chinese life expectancy worsens the U.S. deficit until 2012. However, a more important finding is that the direction of this effect is reversed as the life expectancy gap continues to narrow, and the increased U.S. deficit with China is reduced by the further catch-up of Chinese life expectancy after 2013. It is also found that the relationship between the savings rate and life expectancy in China is reversed for the same period. Furthermore, these results are obtained by using the recently developed technique of polynomial cointegration. Therefore, the U-shaped impact of the life expectancy gap is a statistically valid result.

These empirical findings are new in the literature, and the theoretical foundation is required. This paper develops a simple and analytically tractable OLG model that is useful for having a clear understanding of the mechanism of the U-shaped impact. We show that this impact can be theoretically reproduced by decomposing the effect of foreign life expectancy into three components—savings for retirement, taxes, and elderly productivity—and both the steady-state and transitional analyses provide similar conclusions.

Finally, we briefly remark on policy implications. The United Nations predicts that the life expectancy gap between the United States and China will fall over the next 60 years as shown in Figure 7. This long-run demographic trend will lead to a persistent improvement of the U.S. current account balance with China, with other things being equal, because the life expectancy gap is below its threshold level. Consequently, U.S.–China trade tensions may be partially resolved. For example, the U.S. current account balance with China as a ratio of U.S. GDP was –1.45% in the first quarter of 2019 and was the highest after 2005. This implication is consistent with Ferrero (2010), who shows that most of the permanent component of the U.S. trade balance with the other G7 countries is described by demographic variables. In contrast, the cyclical fluctuations in the U.S. current account balance with China are likely to be affected by other factors such as the business cycle and the real U.S. dollar exchange rate, which are also found to have a significant impact. In particular, the cyclical fluctuations will be large because cyclical factors are invariably much more volatile than demographic factors.

Figure 7 around here

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Appendix A. Data

Data	Source
U.S. current account balance with China, SA	U.S. Bureau of Economic Analysis
Life expectancy at birth	World Development Indicators
GDP, SA	U.S. Bureau of Economic Analysis National Bureau of Statistics of China
Domestic demand (calculated as GDP minus net exports of goods and services), SA	U.S. Bureau of Economic Analysis State Administration of Foreign Exchange
U.S. GDF deflator, SA Consumer price index, SA	U.S. Bureau of Economic Analysis OECD, Main Economic Indicators
China/U.S. foreign exchange rate	Board of Governors of the Federal Reserve System
S&P 500 volatility index	Chicago Board Options Exchange
Wilshire 5000 total market full cap index, SA China's stock market capitalization of stock exchanges, SA	Wilshire Associates Datastream
Population aged 15–64 years, SA	OECD, Main Economic Indicators World Development Indicators
Government net lending/borrowing, SA	U.S. Bureau of Economic Analysis World Economic Outlook Database

Seasonal adjustment is not applicable to the data on life expectancy because the annual data are converted into quarterly data. The basic estimation results remain unchanged if we use annual data. For details, see Sections 3.2 and 4.2.

Appendix B. Test for the relationship between elderly productivity and life expectancy

Our model assumes that better health mitigates a decline in elderly productivity (Eq. 7). As mentioned in Section 7.1, this relationship is a key component of the comparative static analysis. To examine its validity empirically, we use cointegration techniques because the existence of long-run equilibrium relationships is a useful criterion for evaluating the validity of model assumptions (e.g., Corbae and Ouliaris, 1988; Miyao, 1996).

Eq. (7) is rewritten as the following regression model:

$$\ln(\delta_t) = \eta + \gamma \ln(p_t) + v_t \,, \tag{B1}$$

where η is a constant term and v_t is an error term. The structural parameter γ is estimated as a coefficient on $\ln(p_t)$. As for the empirical analyses in Sections 4 and 5, we use the U.S. data for the sample period from the first quarter of 2003 to the fourth quarter of 2015 (the data on life expectancy at age 65 years explained below are available until 2015 in the present study).

From Eqs (8) and (9), the wage rates for young and old people are w_t and $\delta_t w_t$, respectively. Therefore, this paper uses the wage rate for old people relative to that for young people as a measure of δ_t. The data consist of seasonally adjusted weekly nominal earnings for full-time workers aged 35–44 years and 65 years and over, and characteristics of workers are the same for both age groups (all industries and occupations; both sexes; all races; all educational levels). The values of the relative wage rate are less than 1 over the sample period, and this is consistent with the assumption that $0 \leq \delta_t \leq 1$. The estimation results remain unchanged if we use the earnings data for full-time workers aged 45–54 years and the average data for two age groups (35–44, 45–54). Life expectancy at age 65 years is used for consistency with the earnings data. Although this variable is converted into quarterly frequency by using the same interpolation method as explained in Section 3.2, the estimation results are robust if we use the annual data. Both variables are standardized by own initial observations because data units are different. The sources are the U.S. Bureau of Labor Statistics and the National Center for Health Statistics.

We find that the relative wage rate and life expectancy at age 65 years have a unit root and they are cointegrated (Table B1a). Hence, Eq. (B1) is regarded as a long-run equilibrium relationship.

Table B1 around here

The estimation results for Eq. (B1) are indicated in Table B1. The FMOLS method is used to correct for endogeneity and serial correlation biases. We find that γ is positive and significant, and η is not statistically different from zero. These results are consistent with the model assumption, and thus Eq. (B1) is successfully estimated.

Supplementary material (not intended for publication)

B.1. Effect of educational attainment

To control for the effect of educational attainment, we use the share of the labor force with a bachelor's degree and higher in the total labor force for the same age group. The share for the old age group (65 years and over) relative to that for the middle age group (35–44 years) in natural logarithms is added to Eq. (B1) for consistency with the relative wage rate. The sample period is the same as in Appendix B. The source is the U.S. Bureau of Labor Statistics.

The estimation results are reported in Table B2 below. The estimated coefficient on life expectancy is similar to that reported in Table B. Therefore, the robust estimation result for γ is obtained after controlling for educational attainment.

Table B2. Robustness check for the estimate of γ

	Estimate	Standard error
Life expectancy at age 65 years	3.034**	[0.819]
Relative educational attainment	0.166	[0.381]
Constant term	–0.015	[0.052]

The dependent variable is the relative wage rate. The FMOLS method is used. Numbers within parentheses are HAC standard errors. ** indicates significance at the 1% level.

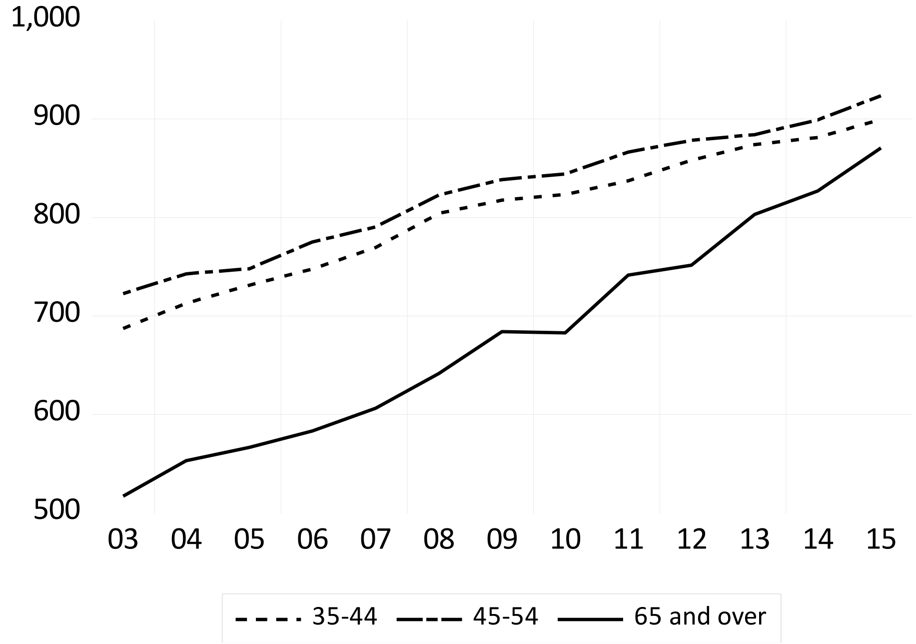
Table B2 also shows that relative educational attainment does not have a significant impact on the relative wage rate ($\delta_t w_t / w_t = \delta_t$). A possible explanation for this result is that educational attainment affects w_t. This means that higher educational levels lead to higher wage rates for both age groups, and thus this effect is excluded from the relative wage rate.

B.2. Wage rate and health status for elderly workers

Figure B1 below shows that the wage rate is lower for elderly than for middle-aged workers, while this gap continues to narrow over the period. The improvement of the wage rate for elderly workers is positively associated with an increase in life expectancy at age 65 years as indicated in Figure B2, which suggests that better health mitigates a decline in elderly productivity. These observations are consistent with the estimation results for Eq. (B1).

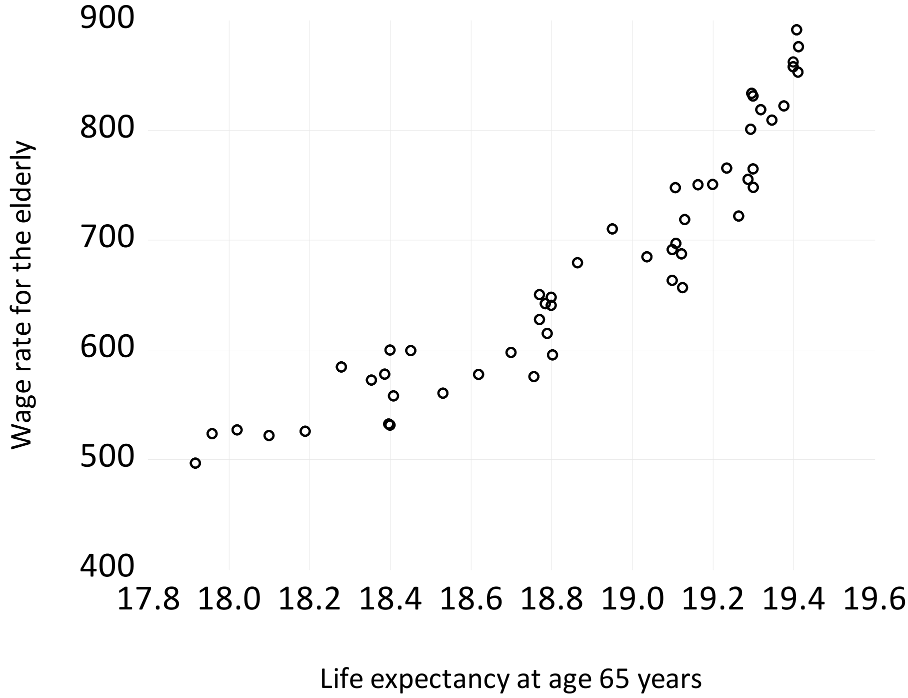
Furthermore, we find that the health status of elderly workers actually improves. The data in Figure B3 are available from 2008 and are standardized by the initial values. The elderly labor force without a disability increased by 1.54 times from 2008 to 2016. Therefore, the number of elderly workers who maintain their health (and thus their productivity) increases over the sample period.

Figure B1. Wage rates for middle-aged and elderly workers



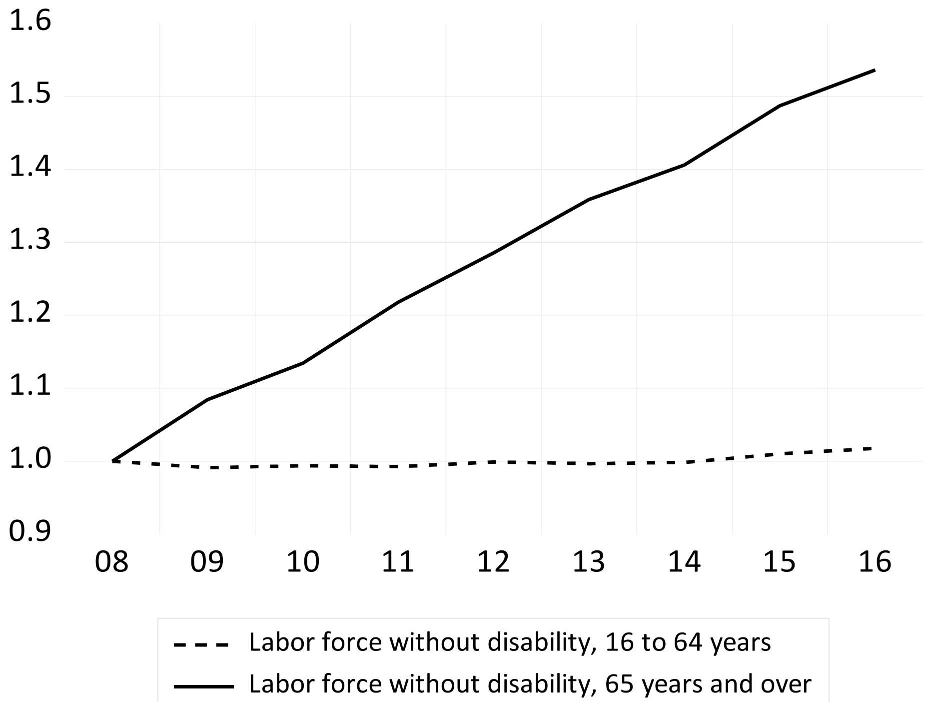
The data consist of weekly nominal earnings for full-time workers (unit: dollar). For all age groups, characteristics of workers are the same (all industries and occupations; both sexes; all races; all educational levels). The source is the U.S. Bureau of Labor Statistics.

Figure B2. The relationship between elderly wage rate and life expectancy



The U.S. quarterly data for the period 2003–2015 are used. The horizontal axis shows life expectancy at age 65 years. The vertical axis shows weekly nominal earnings for full-time workers aged 65 years and over, and the definition is the same as in Figure B1. The sources are the National Center for Health Statistics and the U.S. Bureau of Labor Statistics.

Figure B3. Labor force without a disability



The source is the U.S. Bureau of Labor Statistics.

Appendix C. Proof of Equation (27)

From Eq. (23), the steady-state capital–labor ratio in the closed foreign economy is given by

$$k_c^* = \left(\frac{\lambda^*}{\tilde{\mu}^*}\right)^{\frac{1}{1-\alpha}} \equiv k_c^*(p^*, \tau^*, \delta^*).$$

We obtain the following derivatives:

$$\frac{\partial k_c^*}{\partial p^*} = \frac{1}{(1+p^*)^2}(1-\tau^*)(1+n-\omega^*\delta^* p^{*2})\Omega > 0\,,$$

$$\frac{\partial k_c^*}{\partial \tau^*} = -\frac{p^*}{1+p^*}(1+n+\omega^*\delta^* p^*)\Omega < 0\,,$$

$$\frac{\partial k_c^*}{\partial \delta^*} = -\frac{p^*}{1+p^*}(1-\tau^*)\left[\frac{p^*}{1+p^*}\frac{1-\alpha}{\alpha}(1-\tau^*)\omega^* + \omega^* p^*\right]\Omega < 0\,,$$

where $\Omega \equiv \bar{A}(\lambda^*)^{\frac{\alpha}{1-\alpha}}(\tilde{\mu}^*)^{\frac{\alpha-2}{1-\alpha}}$. In addition, $\frac{d\delta^*}{dp^*} > 0$ and $\frac{d\tau^*}{dp^*} > 0$ hold from the foreign counterparts of Eqs (7) and (14), respectively.

Table 1. Unit root test

	Level	First difference
U.S. current account balance with China	–1.078	–5.024**
Life expectancy gap	3.519	–5.521**
Growth rate difference in real domestic demand	–1.778	–3.037**
Real exchange rate	–0.637	–2.428*
VIX	–1.620	–8.130**
Relative SMC/GDP ratio	–1.459	–3.177**
Growth rate difference in working-age population	–1.197	–4.683**
Difference in fiscal balance/GDP ratio	–0.902	–3.369**

The Dickey–Fuller test with GLS detrending developed by Elliott et al. (1996) is used. The lag length of the regression is selected by the modified Akaike information criterion developed by Ng and Perron (2001). The test includes a constant term. ** and * indicate significance at the 1% and 5% levels, respectively.

Supplementary material (not intended for publication)

If we use the unit root test developed by Phillips and Perron (1988), the results remain unchanged as shown below.

Table 1a. Alternative unit root test: Phillips and Perron (1988)

	Level	First difference
U.S. current account balance with China	–2.634	–5.396**
Life expectancy gap	1.138	–7.214**
Growth rate difference in real domestic demand	–2.319	–7.194**
Real exchange rate	–1.332	–3.552*
VIX	–2.746	–8.377**
Relative SMC/GDP ratio	–2.523	–5.476**
Growth rate difference in working-age population	–1.659	–7.434**
Difference in fiscal balance/GDP ratio	–1.347	–5.473**

Similar to Table 1, the null hypothesis of a unit root is tested. ** and * indicate significance at the 1% and 5% levels, respectively. The test includes a constant term.

Table 2. Test for quadratic polynomial cointegration

Null hypothesis	CT test statistic
The life expectancy gap has a cointegrating quadratic polynomial relationship to the U.S. current account balance with China.	0.085

The long-run variance is estimated by the Bartlett kernel, and the bandwidth parameter is selected by the Newey and West (1987) procedure. Eq. (1) for $m = 1$ is used, and the critical values are tabulated in Wagner (2013).

Supplementary material (not intended for publication)

The growth rate difference in working-age population (another demographic factor) does not have a nonlinear impact on the U.S. current account balance with China. The result is presented below. Therefore, although the old-age dependency ratio is decomposed into life expectancy and the growth rate of working-age population, only the former has a nonlinear impact. This discussion is provided in footnote 14.

Table 2a. Test for quadratic polynomial cointegration: working-age population

Null hypothesis	CT test statistic
The growth rate difference in working-age population has a cointegrating quadratic polynomial relationship to the U.S. current account balance with China.	0.306*

* indicates that the null hypothesis of quadratic polynomial cointegration is rejected at the 5% significance level.

Table 3. Long-run equilibrium equation for the U.S. current account balance with China

	FMOLS	DOLS
(A) Basic model		
Life expectancy gap (US – China)	–6.915**	–5.658**
	[0.954]	[1.122]
Squared life expectancy gap	1.137**	0.941**
	[0.149]	[0.173]
Constant term	8.454**	6.432**
	[1.513]	[1.829]
Threshold of life expectancy gap	3.040	3.005
(B) Full model		
Life expectancy gap (US – China)	–7.032**	–8.108**
	[0.717]	[0.760]
Squared life expectancy gap	1.208**	1.420**
	[0.116]	[0.123]
Growth rate difference in real domestic demand (US – China)	–0.018**	–0.038**
	[0.006]	[0.006]
Real exchange rate	–0.899**	–1.874**
	[0.323]	[0.247]
VIX	0.225*	0.347**
	[0.065]	[0.077]
Relative SMC/GDP ratio (US/China)	0.269**	0.330**
	[0.057]	[0.070]
Growth rate difference in working-age population (US – China)	0.221**	0.170*
	[0.070]	[0.072]
Difference in fiscal balance/GDP ratio (US – China)	0.010	0.020**
	[0.009]	[0.006]
Constant term	10.987**	13.993**
	[1.468]	[1.640]
Threshold of life expectancy gap	2.911	2.855

(Continued on the next page)

Table 3 (continued)

	FMOLS	DOLS
(C) Basic model (annual data)		
Life expectancy gap (US – China)	–5.658**	–3.922**
	[0.958]	[0.480]
Squared life expectancy gap	0.951**	0.666**
	[0.154]	[0.079]
Constant term	6.365**	3.700**
	[1.478]	[0.726]
Threshold of life expectancy gap	2.976	2.944

Numbers within parentheses are heteroskedasticity and autocorrelation consistent (HAC) standard errors. The long-run variance is estimated by the Bartlett kernel, and the bandwidth parameter is selected by the Newey and West (1987) procedure. ** and * indicate significance at the 1% and 5% levels, respectively.

Table 4. Out-of-sample forecast performance of linear and nonlinear models

	RMSE	
Specification of the impact of life expectancy gap	Basic model	Full model
Linear	0.620	0.601
Quadratic polynomial	0.219	0.157

The estimation period is from the first quarter of 2003 to the fourth quarter of 2015, and the FMOLS method is used. The forecast period is from the first quarter of 2016 to the fourth quarter of 2016.

Supplementary material (not intended for publication)

Table 4a. Estimation results for Eq. (1): subsample period (2003:Q1–2015:Q4)

	Basic model	Full model
Life expectancy gap (US – China)	–8.124**	–8.313**
	[1.441]	[1.258]
Squared life expectancy gap	1.312**	1.383**
	[0.217]	[0.186]
Growth rate difference in real domestic demand (US – China)		–0.021**
		[0.006]
Real exchange rate		–0.773*
		[0.372]
VIX		0.172*
		[0.073]
Relative SMC/GDP ratio (US/China)		0.281**
		[0.063]
Growth rate difference in working-age population (US – China)		0.192*
		[0.075]
Difference in fiscal balance/GDP ratio (US – China)		0.003
		[0.010]
Constant term	10.530**	13.213**
	[2.378]	[2.221]
Threshold of life expectancy gap	3.096	3.005

If we use the data for the subsample period, the estimation results for Eq. (1) reported in Table 3 remain essentially unchanged. The results of the FMOLS estimation are presented in Table 4a. Therefore, the forecast analysis is conducted using the estimation results consistent with those for the full sample period. For the estimation procedures, see note of Table 3. ** and * indicate significance at the 1% and 5% levels, respectively.

In addition, the results for the forecast evaluation remain essentially unchanged if we use alternative criteria such as the mean absolute percentage error and the Theil inequality coefficient. The results are shown below.

Table 4b. Forecast evaluation based on alternative criteria

	Mean absolute percentage error	
Specification of the impact of life expectancy gap	Basic model	Full model
Linear	34.943	33.005
Quadratic polynomial	12.263	8.400

	Theil inequality coefficient	
Specification of the impact of life expectancy gap	Basic model	Full model
Linear	0.149	0.146
Quadratic polynomial	0.066	0.046

Table 5. Robustness checks for the U-shaped impact of the life expectancy gap

	Nonlinear impact of life expectancy gap (x_1)		
Control variables	x_1	x_1^2	Threshold of x_1
x_2	–6.620** [1.074]	1.091** [0.166]	3.033
x_3	–6.945** [0.852]	1.178** [0.141]	2.947
x_4	–6.476** [0.817]	1.067** [0.128]	3.035
x_5	–6.585** [0.898]	1.086** [0.141]	3.031
x_6	–6.788** [0.950]	1.133** [0.150]	2.997
x_7	–7.423** [1.072]	1.217** [0.167]	3.051
x_2, x_3	–6.925** [1.056]	1.171** [0.172]	2.955
x_2, x_4	–6.231** [0.883]	1.031** [0.137]	3.021
x_2, x_5	–6.544** [1.023]	1.072** [0.158]	3.052
x_2, x_6	–6.605** [1.063]	1.101** [0.166]	2.999
x_2, x_7	–7.066** [1.012]	1.157** [0.157]	3.053
x_3, x_4	–6.681** [0.786]	1.127** [0.129]	2.963
x_3, x_5	–7.141** [0.871]	1.209** [0.143]	2.952
x_3, x_6	–6.642** [0.853]	1.136** [0.141]	2.924
x_3, x_7	–6.871** [0.946]	1.165** [0.150]	2.948
x_4, x_5	–6.387** [0.740]	1.051** [0.116]	3.039
x_4, x_6	–6.210** [0.808]	1.035** [0.128]	3.000
x_4, x_7	–6.760** [1.045]	1.112** [0.163]	3.040
x_5, x_6	–6.442** [0.889]	1.081** [0.140]	2.980
x_5, x_7	–7.622** [1.051]	1.245** [0.163]	3.061
x_6, x_7	–7.185** [1.068]	1.191** [0.167]	3.016
x_2, x_3, x_4	–6.737** [0.918]	1.136** [0.150]	2.964
x_2, x_3, x_5	–8.118** [1.063]	1.362** [0.173]	2.980
x_2, x_3, x_6	–6.647** [0.998]	1.133** [0.163]	2.934
x_2, x_3, x_7	–6.876** [0.973]	1.165** [0.157]	2.951
x_2, x_4, x_5	–6.381** [0.848]	1.046** [0.131]	3.051
x_2, x_4, x_6	–6.097** [0.840]	1.022** [0.132]	2.984
x_2, x_4, x_7	–6.344** [0.790]	1.044** [0.122]	3.039
x_2, x_5, x_6	–6.588** [0.990]	1.098** [0.155]	3.001
x_2, x_5, x_7	–7.406** [1.070]	1.204** [0.165]	3.074
x_2, x_6, x_7	–6.876** [0.922]	1.136** [0.144]	3.027

(Continued on the next page)

Table 5 (continued)

Control variables	Nonlinear impact of life expectancy gap (x_1)		
	x_1	x_1^2	Threshold of x_1
x_3, x_4, x_5	–7.003** [0.742]	1.177** [0.122]	2.976
x_3, x_4, x_6	–6.227** [0.784]	1.058** [0.129]	2.944
x_3, x_4, x_7	–6.308** [0.903]	1.078** [0.142]	2.925
x_3, x_5, x_6	–6.760** [0.871]	1.156** [0.143]	2.924
x_3, x_5, x_7	–7.395** [0.982]	1.247** [0.155]	2.966
x_3, x_6, x_7	–6.611** [0.936]	1.128** [0.149]	2.930
x_4, x_5, x_6	–6.005** [0.664]	1.011** [0.105]	2.971
x_4, x_5, x_7	–6.478** [0.971]	1.137** [0.151]	3.055
x_4, x_6, x_7	–6.446** [1.026]	1.074** [0.160]	3.000
x_5, x_6, x_7	–7.455** [1.001]	1.235** [0.157]	3.017
x_2, x_3, x_4, x_5	–7.787** [0.914]	1.300** [0.149]	2.995
x_2, x_3, x_4, x_6	–6.477** [0.850]	1.099** [0.139]	2.946
x_2, x_3, x_4, x_7	–6.248** [0.848]	1.064** [0.136]	2.936
x_2, x_3, x_5, x_6	–7.674** [0.947]	1.305** [0.156]	2.941
x_2, x_3, x_5, x_7	–7.916** [1.041]	1.331** [0.167]	2.973
x_2, x_3, x_6, x_7	–6.460** [0.889]	1.099** [0.144]	2.934
x_2, x_4, x_5, x_6	–6.256** [0.723]	1.046** [0.113]	2.991
x_2, x_4, x_5, x_7	–6.641** [0.928]	1.083** [0.143]	3.065
x_2, x_4, x_6, x_7	–6.171** [0.928]	1.027** [0.144]	3.005
x_2, x_5, x_6, x_7	–7.558** [0.940]	1.247** [0.147]	3.031
x_3, x_4, x_5, x_6	–6.281** [0.676]	1.070** [0.111]	2.936
x_3, x_4, x_5, x_7	–6.810** [0.885]	1.154** [0.140]	2.952
x_3, x_4, x_6, x_7	–6.018** [0.886]	1.035** [0.140]	2.906
x_3, x_5, x_6, x_7	–7.031** [0.919]	1.191** [0.146]	2.951
x_4, x_5, x_6, x_7	–6.568** [0.837]	1.097** [0.130]	2.994
x_2, x_3, x_4, x_5, x_6	–7.126** [0.715]	1.207** [0.117]	2.953
x_2, x_3, x_4, x_5, x_7	–7.249** [0.876]	1.224** [0.140]	2.962
x_2, x_3, x_4, x_6, x_7	–5.846** [0.900]	1.006** [0.145]	2.907
x_2, x_3, x_5, x_6, x_7	–7.707** [0.901]	1.304** [0.146]	2.950
x_2, x_4, x_5, x_6, x_7	–6.618** [0.684]	1.108** [0.106]	2.987
x_3, x_4, x_5, x_6, x_7	–6.319** [0.764]	1.078** [0.121]	2.931

x_2 is the growth rate difference in real domestic demand; x_3, the real exchange rate; x_4, VIX; x_5, the relative SMC/GDP ratio; x_6, the growth rate difference in working-age population; and x_7, the difference in the fiscal balance/GDP ratio. The FMOLS method is used. Numbers within parentheses are HAC standard errors. ** indicates significance at the 1% level.

Supplementary material (not intended for publication)

Table 5a. Robustness checks based on the DOLS estimation

	Nonlinear impact of life expectancy gap (x_1)		
Control variables	x_1	x_1^2	Threshold of x_1
x_2	–5.549** [1.061]	0.928** [0.163]	2.990
x_3	–5.402** [0.996]	0.945** [0.158]	2.859
x_4	–4.940** [0.865]	0.830** [0.133]	2.976
x_5	–5.280** [0.987]	0.874** [0.152]	3.021
x_6	–5.579** [0.763]	0.943** [0.181]	2.957
x_7	–6.118** [0.962]	1.016** [0.149]	3.010
x_2, x_3	–5.725** [0.879]	1.005** [0.141]	2.848
x_2, x_4	–4.814** [0.797]	0.814** [0.122]	2.957
x_2, x_5	–5.105** [1.060]	0.851** [0.161]	2.999
x_2, x_6	–5.437** [1.085]	0.920** [0.171]	2.955
x_2, x_7	–6.202** [0.986]	1.029** [0.152]	3.013
x_3, x_4	–5.164** [0.835]	0.898** [0.133]	2.874
x_3, x_5	–5.155** [1.078]	0.900** [0.173]	2.865
x_3, x_6	–3.984** [0.948]	0.745** [0.149]	2.672
x_3, x_7	–5.498** [0.839]	0.954** [0.130]	2.883
x_4, x_5	–4.905** [0.855]	0.819** [0.131]	2.995
x_4, x_6	–4.760** [0.841]	0.805** [0.133]	2.956
x_4, x_7	–4.601** [1.016]	0.781** [0.157]	2.945
x_5, x_6	–5.219** [0.958]	0.886** [0.150]	2.946
x_5, x_7	–5.593** [1.225]	0.930** [0.189]	3.006
x_6, x_7	–5.951** [0.929]	0.985** [0.147]	3.021
x_2, x_3, x_4	–5.643** [0.870]	0.984** [0.142]	2.868
x_2, x_3, x_5	–7.862** [1.659]	1.355** [0.273]	2.901
x_2, x_3, x_6	–4.411** [0.934]	0.824** [0.148]	2.676
x_2, x_3, x_7	–5.842** [0.719]	1.053** [0.114]	2.774
x_2, x_4, x_5	–4.817** [0.859]	0.810** [0.130]	2.973
x_2, x_4, x_6	–4.582** [0.747]	0.779** [0.118]	2.940
x_2, x_4, x_7	–4.558** [1.005]	0.775** [0.155]	2.941
x_2, x_5, x_6	–5.227** [1.060]	0.889** [0.164]	2.940
x_2, x_5, x_7	–5.692** [1.362]	0.945** [0.208]	3.012
x_2, x_6, x_7	–6.004** [0.946]	0.995** [0.150]	3.016

(Continued on the next page)

Table 5a (continued)

Control variables	Nonlinear impact of life expectancy gap (x_1)		
	x_1	x_1^2	Threshold of x_1
x_3, x_4, x_5	–4.976** [1.144]	0.873** [0.177]	2.849
x_3, x_4, x_6	–3.844** [0.720]	0.704** [0.112]	2.729
x_3, x_4, x_7	–4.343** [0.828]	0.781** [0.128]	2.782
x_3, x_5, x_6	–4.138** [1.006]	0.765** [0.158]	2.704
x_3, x_5, x_7	–3.763** [1.254]	0.693** [0.192]	2.716
x_3, x_6, x_7	–4.211** [0.896]	0.763** [0.138]	2.759
x_4, x_5, x_6	–5.360** [0.721]	0.913** [0.116]	2.935
x_4, x_5, x_7	–4.075** [1.175]	0.696** [0.181]	2.929
x_4, x_6, x_7	–4.247** [0.946]	0.722** [0.148]	2.940
x_5, x_6, x_7	–6.002** [1.265]	1.000** [0.198]	3.002
x_2, x_3, x_4, x_5	–8.174** [1.377]	1.401** [0.222]	2.917
x_2, x_3, x_4, x_6	–4.439** [0.876]	0.806** [0.140]	2.754
x_2, x_3, x_4, x_7	–4.861** [0.589]	0.896** [0.094]	2.713
x_2, x_3, x_5, x_6	–7.562** [0.974]	1.352** [0.161]	2.796
x_2, x_3, x_5, x_7	–6.563** [1.295]	1.177** [0.208]	2.788
x_2, x_3, x_6, x_7	–4.881** [0.782]	0.902** [0.123]	2.706
x_2, x_4, x_5, x_6	–5.879** [0.789]	1.000** [0.127]	2.939
x_2, x_4, x_5, x_7	–4.136** [1.230]	0.706** [0.188]	2.978
x_2, x_4, x_6, x_7	–4.195** [0.920]	0.715** [0.144]	2.934
x_2, x_5, x_6, x_7	–6.533** [1.500]	1.084** [0.234]	3.015
x_3, x_4, x_5, x_6	–4.589** [0.975]	0.822** [0.152]	2.791
x_3, x_4, x_5, x_7	–3.502** [1.206]	0.651** [0.184]	2.688
x_3, x_4, x_6, x_7	–3.451** [0.842	0.641** [0.130]	2.691
x_3, x_5, x_6, x_7	–4.064** [1.190]	0.742** [0.184]	2.740
x_4, x_5, x_6, x_7	–4.760** [1.027]	0.815** [0.161]	2.919
x_2, x_3, x_4, x_5, x_6	–8.697** [0.814]	1.513** [0.133]	2.874
x_2, x_3, x_4, x_5, x_7	–7.059** [0.977]	1.250** [0.156]	2.823
x_2, x_3, x_4, x_6, x_7	–4.322** [0.672]	0.802** [0.107]	2.695
x_2, x_3, x_5, x_6, x_7	–7.232** [0.947]	1.291** [0.153]	2.801
x_2, x_4, x_5, x_6, x_7	–5.817** [1.143]	0.984** [0.180]	2.954
x_3, x_4, x_5, x_6, x_7	–3.915** [1.147]	0.712** [0.177]	2.748

The number of leads and lags in the DOLS regression is 1. Numbers within parentheses are HAC standard errors. The definitions of the control variables are the same as in Table 5. ** indicates significance at the 1% level.

Table 6. Relationship between the savings rate and life expectancy in China

	Coefficient	Standard error
Life expectancy	257.752**	[7.335]
Squared life expectancy	–1.704**	[0.049]
Constant term	–9705.231**	[275.006]
Threshold of life expectancy	75.631	

The estimation method is the FMOLS. Numbers within parentheses are HAC standard errors. ** indicates significance at the 1% level.

Table 7. Estimation of error correction model

	$m = 1$	$m = 7$
(A) Estimation results		
Error correction term	–0.244**	–0.311*
	[0.080]	[0.117]
Constant term	–0.012	–0.014
	[0.017]	[0.012]
<u>Lagged variables in first differences</u>		
U.S. current account balance with China	0.452**	0.422**
	[0.127]	[0.128]
Life expectancy gap (US – China)	–0.226	–0.379
	[0.189]	[0.247]
Growth rate difference in real domestic demand (US – China)		0.002
		[0.004]
Real exchange rate		0.653
		[0.663]
VIX		0.013
		[0.043]
Relative SMC/GDP ratio (US/China)		–0.163**
		[0.058]
Growth rate difference in working-age population (US – China)		–0.006
		[0.059]
Difference in fiscal balance/GDP ratio (US – China)		–0.007
		[0.011]
(B) Residual diagnostics		
Jarque–Bera statistic for the null hypothesis of normality	3.396	1.214
	(0.183)	(0.545)
Q-statistic for the null hypothesis of no residual autocorrelation	4.201	2.352
	(0.380)	(0.671)

m is the number of integrated regressors, and the specifications for $m = 1$ and 7 are based on the basic and full models used in the cointegration analysis, respectively. The error correction term is calculated as the FMOLS residual in Eq. (1), and the coefficients reported in Table 3 are used. The null hypothesis of no residual autocorrelation up to order 4 is tested. Numbers within square brackets and round parentheses are standard errors and p-values, respectively. ** and * indicate significance at the 1% and 5% levels, respectively.

Table 8. Parameters for numerical analysis

Parameter	Description	Value
$\alpha = \alpha^*$	Capital share of output	0.409
θ	Population share of the home country	0.500
$\bar{A} = \bar{A}^*$	Technology parameter	10.000
$n = n^*$	Growth rate of the population aged 15–64 years	0.365
$\omega = \omega^*$	Labor force participation rate of people aged 65 years and over	0.202
$\phi = \phi^*$	Net replacement rate	0.644
$\gamma = \gamma^*$	Author's estimation (Appendix B)	3.195

To calculate the capital share, the average of the U.S. labor income share $(1 - \alpha)$ for the period 2000–2016 is used, and the source is Giandrea and Sprague (2017). The values of population share and technology parameter are the same as those used in Ito and Tabata (2010). The average of the annual growth rates of the population aged 15–64 for the United States and China over the period 2003–2016 is 0.0078, and the parameter for simulation is calculated as $1.0078^{40} - 1$. The source is the World Development Indicators of the World Bank. The labor force participation rate of people aged 65 years and over is calculated from the average of the data for the United States in 2016 and China in 2010 because of data availability. Similarly, the average of the data on the net replacement rate (male) for the United States and China in 2014 and 2016 is used. The source of these data is the OECD.

Table 9. Numerical comparative statics: Effect of p^* on k_c^* based on Eq. (27)

p^*	0.80	0.81	0.82	0.83	0.84
Saving effect	1.136	1.108	1.080	1.053	1.025
Tax effect	–0.387	–0.379	–0.371	–0.362	–0.353
Productivity effect	–0.646	–0.668	–0.690	–0.711	–0.733
Total effect	0.103	0.061	0.020	–0.021	–0.061

The parameters reported in Table 8 are used. The initial condition is given by $p^* = 0.80$, and then only p^* is increased to describe the catching-up process of foreign life expectancy.

Table B1. Relationship between wage rate and life expectancy

	Coefficient	Standard error
Life expectancy at age 65 years	3.195**	[0.500]
Constant term	–0.028	[0.029]

The FMOLS method is used. Numbers within parentheses are HAC standard errors. ** indicates significance at the 1% level.

Supplementary material (not intended for publication)

Table B1a. Unit root and cointegration tests for Eq. (B1)

(A) Unit root test (Elliott et al., 1996)		
	Level	First difference
Relative wage rate	–0.130	–4.032**
Life expectancy at age 65 years	1.441	–2.142*

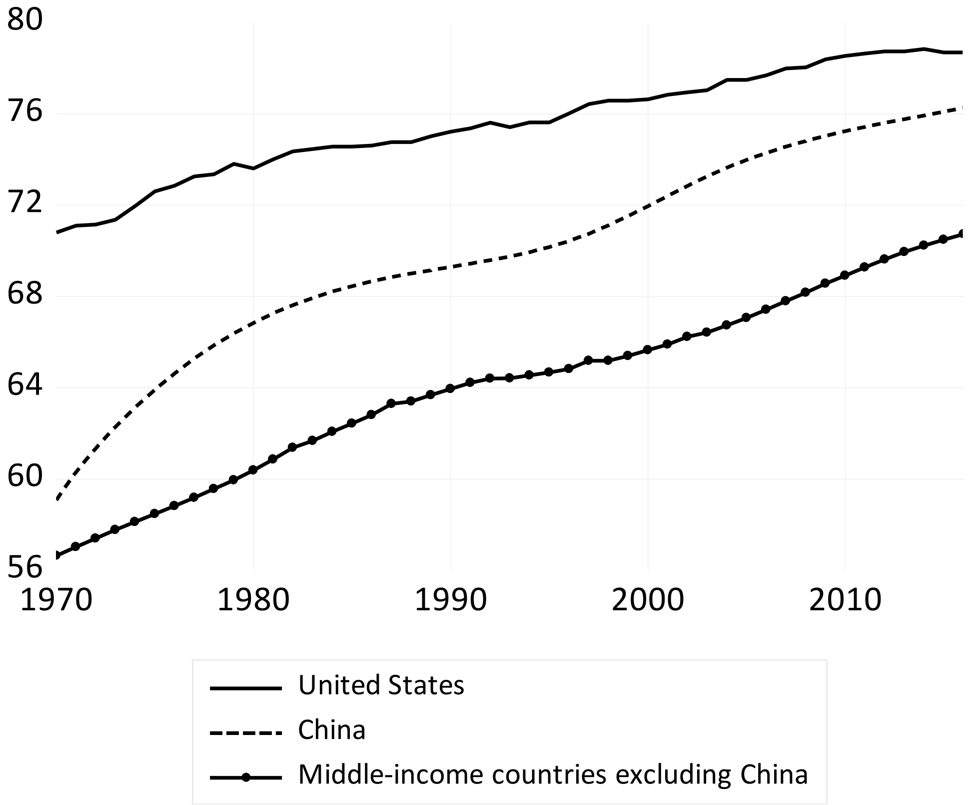
(B) Linear cointegration test (Engle and Granger, 1987)	
Test statistic	–3.983*

** and * indicate significance at the 1% and 5% levels, respectively.

The unit root test does not reject the null hypothesis of a unit root for the variables in levels. However, the null hypothesis of a unit root is rejected for the variables in first differences. These results suggest that both variables are integrated of order one.

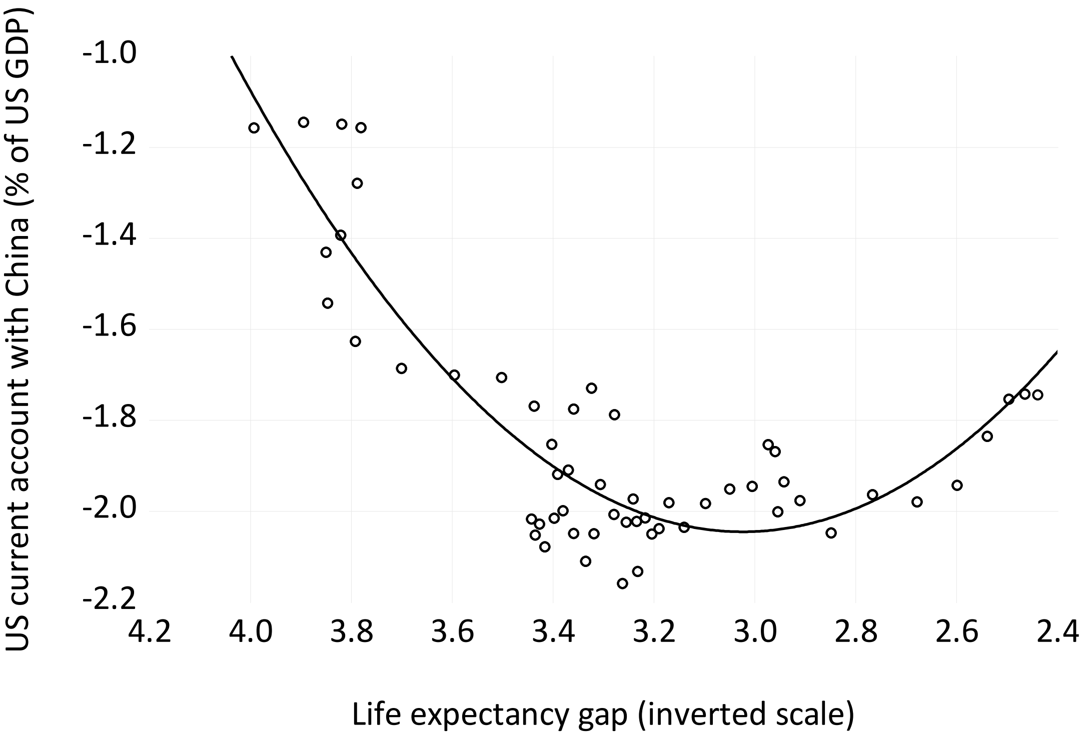
The null hypothesis of no linear cointegration is rejected at the 5% significance level. Therefore, the relative wage rate and life expectancy at age 65 years are cointegrated, and Eq. (B1) is estimated as a linear long-run equilibrium relationship.

Figure 1. Life expectancy at birth



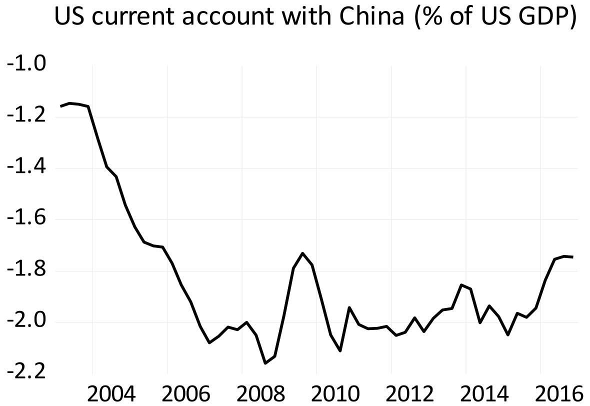
Middle-income countries are those with a GNI per capita between $996 and $12,055, and this definition is based on the country classification of the World Bank for the 2019 fiscal year. Life expectancy for the middle-income group is calculated as a cross-country average for every year. The source is the World Development Indicators of the World Bank.

Figure 2. U-shaped impact of the life expectancy gap on the U.S. deficit with China



The sample period is from the first quarter of 2003 to the fourth quarter of 2016 because of data availability. Detailed explanations for the data are provided in Section 3.2 and Appendix A. The life expectancy gap is defined as the subtraction of Chinese life expectancy at birth from U.S. life expectancy at birth, and a narrowing life expectancy gap means the catching-up of Chinese life expectancy. The nonlinear fitted line is derived from the estimation results of the polynomial cointegrating regression model reported in Section 4.2.

Supplementary material (not intended for publication)



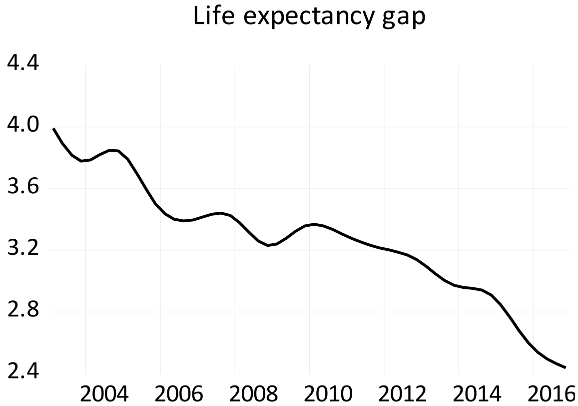
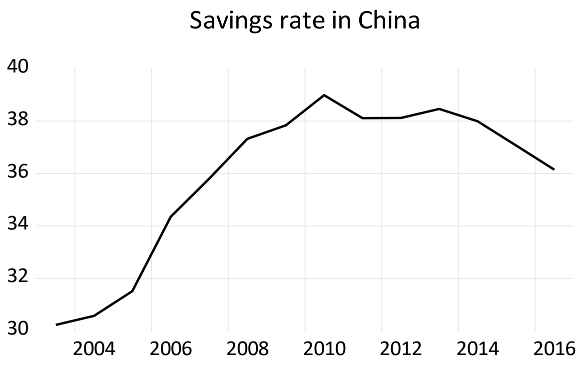


Figure 3. Nonlinear relationship between the savings rate and life expectancy in China

The nonlinear fitted line is derived from the estimation results of the polynomial cointegrating regression model reported in Table 6.

Supplementary material (not intended for publication)



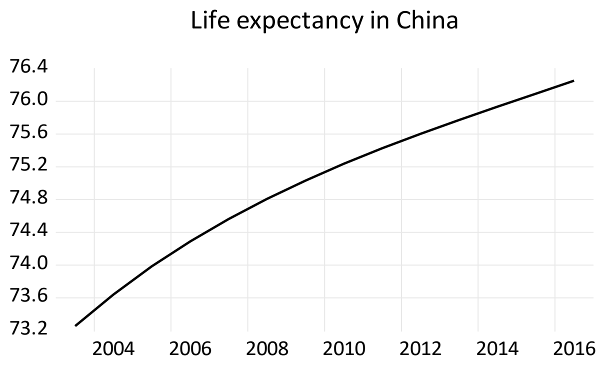
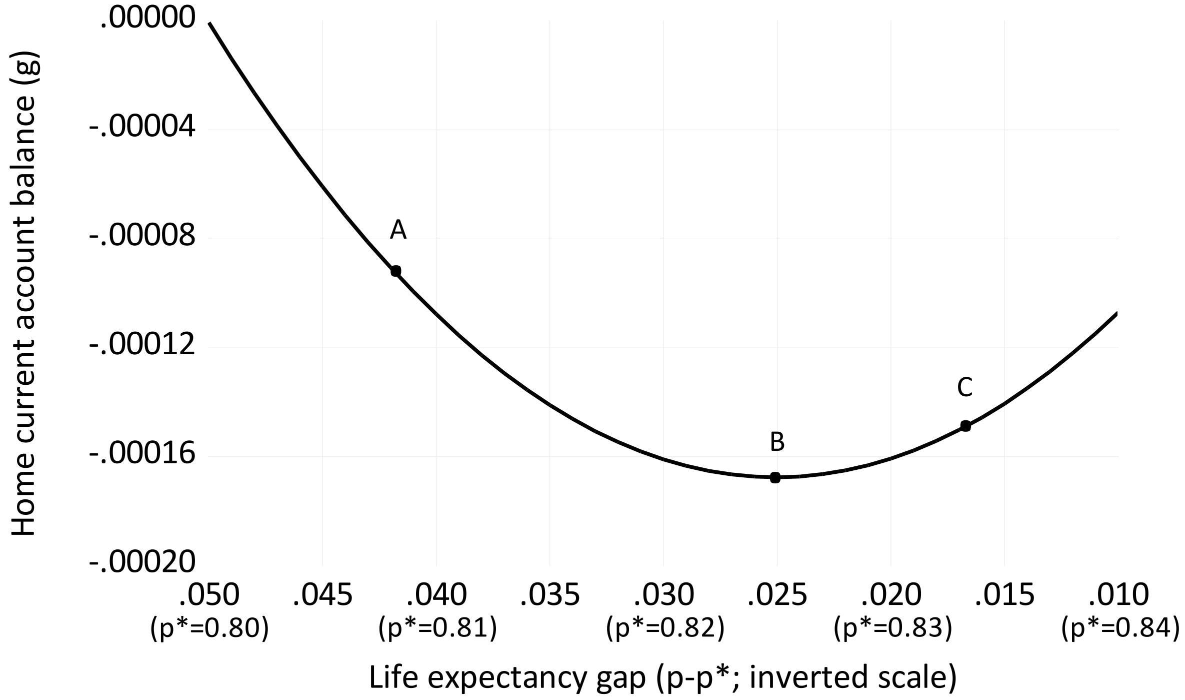
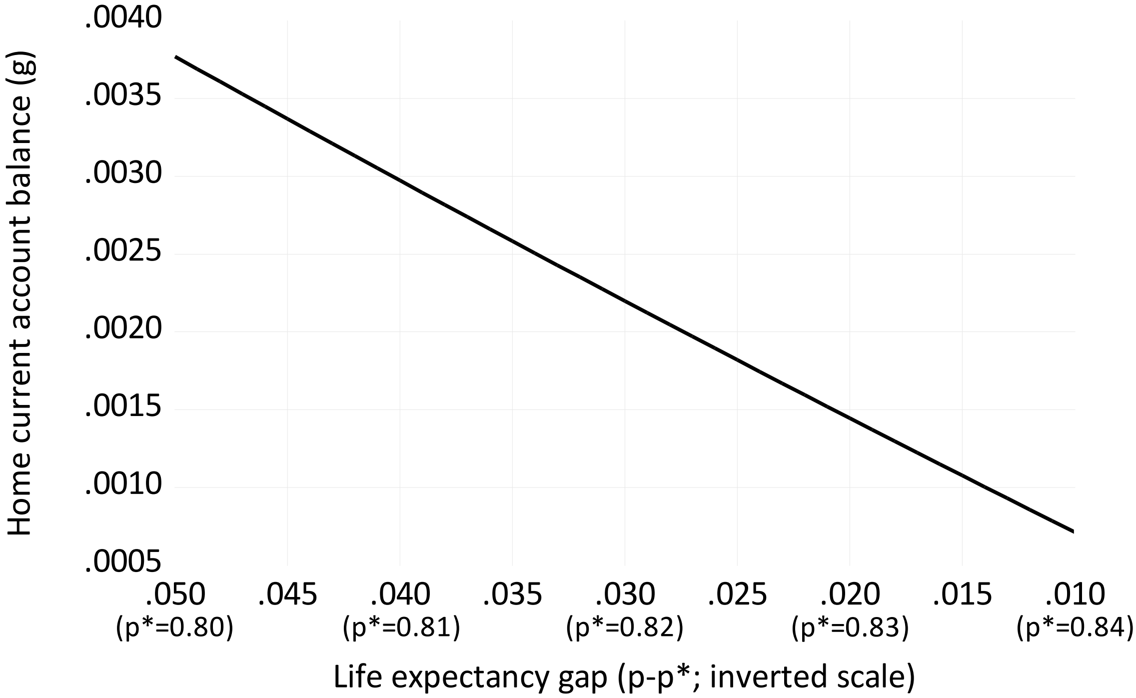


Figure 4. Catch-up of foreign life expectancy and its impact on the home country



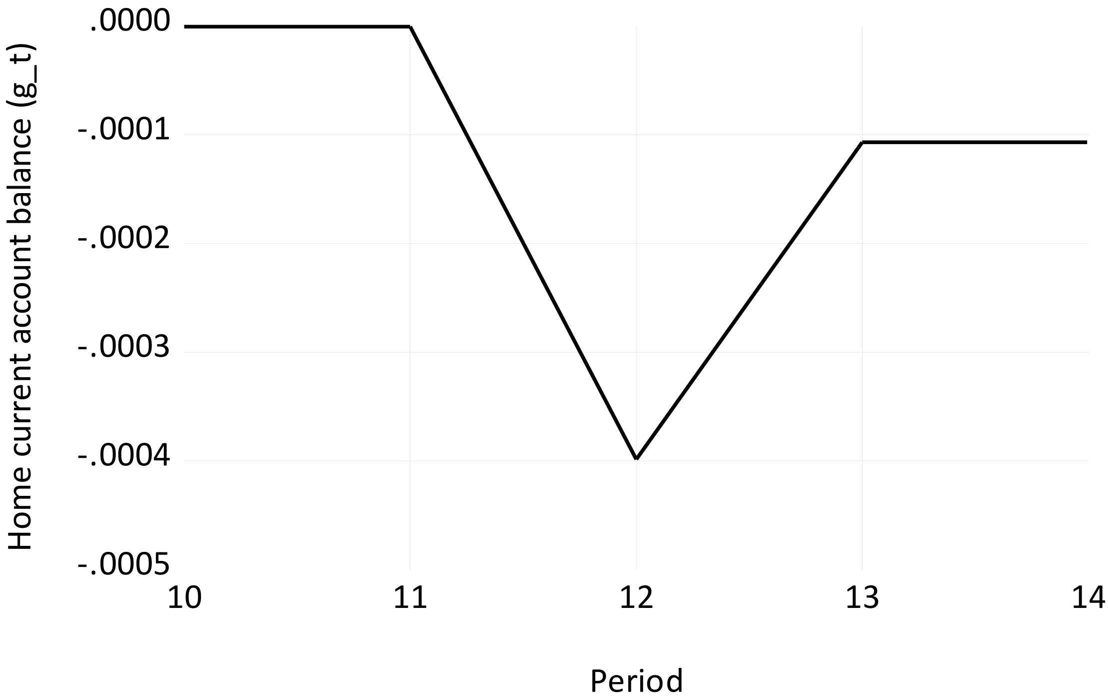
The parameters reported in Table 8 are used. The initial conditions are given by $p = 0.85$ and $p^* = 0.80$, and then only p^* is increased.

Figure 5. Relationship between g and $p - p^*$ under the assumption $\omega = \omega^* = 0$



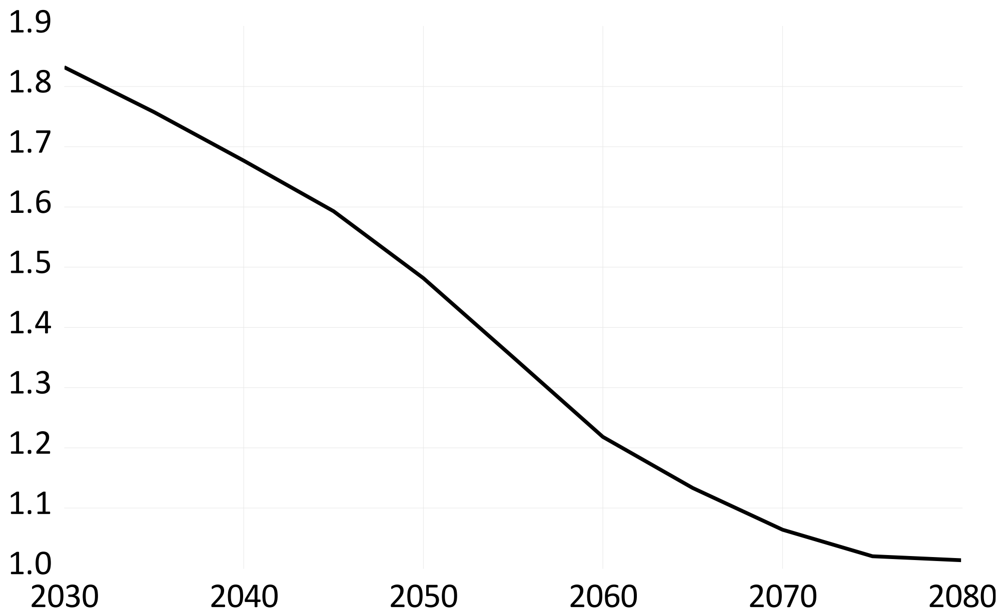
The labor force participation rates of the elderly in the home and foreign countries (ω and ω^*) are assumed to be zero. The other parameters are the same as in Table 8. The initial conditions are given by $p = 0.85$ and $p^* = 0.80$, and then only p^* is increased.

Figure 6. Transition dynamics of the home current account balance



The parameters reported in Table 8 are used. For $t < 11$, the world economy is in a steady-state equilibrium where $p = 0.85$ and $p^* = 0.80$. For $t \geq 11$, p^* is increased to 0.84.

Figure 7. Prospect of the life expectancy gap between the United States and China



The forecast data on U.S. and Chinese life expectancy at birth after 2030 are obtained from the World Population Prospects of the United Nations (the 2019 Revision; medium variant). The life expectancy gap is calculated as the subtraction of Chinese life expectancy at birth from U.S. life expectancy at birth.