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# Can we have growth when population is stagnant? Testing linear growth rate formulas and their cross-unit cointegration of non-scale endogenous growth models<sup>1</sup>

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*Abstract* We sub-divide scale-invariant fully or semi-endogenous growth models into six sub-categories for formulas relating steady-state growth rates of income per capita and the growth rate of the population depending on the properties of slopes and intercepts. We capture their steady-state relation by a long-term relation in panel vector-error-correction models for 16 countries, and estimate the 16 models simultaneously allowing successively for more heterogeneity. The slope and intercepts of the growth equations are positive in this setting under slope homogeneity but less significant or even negative when allowing for heterogeneity. Slopes are mostly non-positive. Intercepts are positive for a large majority of countries. Results therefore favour fully over semi-endogenous growth with and without slope homogeneity and allow for growth rate policies. The more frequent case is that long-run growth can remain positive if population stops growing. Analysis of cross-unit cointegration suggests that long-run results are internationally connected.

JEL-codes: C33, O47. Keywords: Endogenous growth, population growth, panel times series estimation.

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<sup>1</sup> This paper extends WP 2018-044 to include cross-unit cointegration in section 3.10.

## 1 Introduction

The literature characterises models of economic growth as semi-endogenous if endogenous technical change goes to zero in case population growth does so because of decreasing returns in innovation production. They were invented by Arrow (1962) and Phelps (1966), extended to micro-foundation by Judd (1985) and to spillovers by Jones (1995) and others. Ever since the invention, a major criticism has been exactly its defining property, that it would predict having no long-run growth in case of the possible long-run absence of population growth (von Weizsäcker 1969). In contrast, Jones (1995) emphasises that the number of engineers and scientists in R&D has increased strongly in the USA during recent years but the growth rate has not reacted and therefore semi-endogenous growth models should be favoured over fully endogenous growth models. The latter are defined as models with positive long-term growth even in the absence of population growth. However, more recent evaluations of endogenous growth models favour the fully endogenous non-scale growth models, defined as having no scale effects from the *level* of labour resource endowments. They emphasise the empirical perspective of the time-series dimension, using econometric unit root and cointegration analysis as well as forecasts (see Ha and Howitt 2007, Madsen 2008, and many references in Barcenilla-Visús et al. 2014 and Peretto 2018). Ha and Howitt (2007) emphasise that R&D as a share of GDP in the USA has been constant and so have the TFP growth rates. Similarly, from the perspective of Lucas' (1988) endogenous growth model, the education time share data of Gaessler and Ziesemer (2016) for the USA are not increasing for the period 1985-2009, and therefore growth rates are constant. Therefore, Jones' (1995) critique, based on data for engineers, would not hold when

using the education data in the Lucas model<sup>2</sup> or the R&D/GDP data of Ha and Howitt (2007)<sup>3</sup>. Moreover, in the closed and the open economy versions of the Lucas model (Lucas 1988, Frenkel et al 1996; Gaessler and Ziesemer 2016) and in some monopolistic competition models, population growth and that of GDP per capita are positively related with a positive intercept in the solution of the model. The latter implies positive growth rates in the absence of population growth. The intercept then may represent institutional aspects, which ensure positive rates of growth even in the absence of resource growth (Ha and Howitt 2007). In Lucas' (1988) model this would be the productivity of the education system.<sup>4</sup> In Howitt (1999) this is the productivity of the R&D process and spillovers. In Dalgaard and Kreiner (2001) this are the elasticities of production of human capital and intermediates, and positive horizontal spillover effects, relative to a dilution of human capital through a higher degree of difficulty in output production, as well as the subsidies to education and R&D. In contrast, some other endogenous growth models have a negative or zero slope, or a zero or even negative intercept.<sup>5</sup> Prettner and Prskawetz (2010) point out that some authors favour the semi-endogenous type of models with positive slope and zero intercept more recently. Studies without panel heterogeneity put a bit more weight on the semi-endogenous growth models (see Neves and Sequeira 2017). This view gets some recent support from Barcenilla-Visús et al. (2014) and Kruse-Andersen

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<sup>2</sup> Even if education time-shares would increase, one would have to check for an increase of the depreciation rate for human capital during the same period in order to see the net effect on growth. Similarly, in the monopolistic competition models rates of depreciation for knowledge are missing which makes empirical work somewhat imprecise.

<sup>3</sup> A constant R&D output/input ratio does not require that stock variables are irrelevant in the production function, but the latter is sufficient for the former (Madsen 2007).

<sup>4</sup> The maximum growth rate of human capital diminished by the rate of time preference if the intertemporal elasticity of marginal utility is unity; for deviations from unity, externalities may have a positive or negative effect.

<sup>5</sup> An exception is Strulik (2005) where the intercept is positive and the slope can have any sign. He combines the Lucas model with that of varieties, including duplication and difficulty. Strulik's (2005) and also Bucci's (2008) synthesis of Lucas' and the variety models can be reconciled with more than one of these outcomes.

(2017) for R&D models estimating R&D functions. Others doubt the positive slope though. Strulik et al. (2013) argue that ‘there is little empirical support for a positive association between population growth and productivity growth’. Similarly, Chang et al. (2017) conclude that the lack of Granger causality between population growth rates and income growth rates makes sense because technological change drives growth. However, it is exactly in the theory of endogenous technological change that the link between these two growth rates is prominent. All these stark contrasts in the views on slopes and intercepts raise two questions, (i) that of a positive, zero or even negative slope relating population growth and income growth, and (ii) that of the (non-) existence of a statistically significant intercept. With evidence favouring fully and semi-endogenous growth mixed, one may expect results that are in the neighbourhood of both, different from sector to sector as in Venturini (2012), and from country to country in this paper, in line with current emphasis on heterogeneity in econometric dynamic panel data analysis.

## **2 Long-run growth formulas under monopolistic competition**

Monopolistic competition models are slightly more complicated than the perfectly competitive Lucas model mentioned above. The semi-endogenous growth model generates the long-run growth formula for per-capita income growth  $g_y = (\lambda - \nu)g_n / (1 - \gamma)$  with  $g_n$  as population growth rate,  $\gamma < 1$  as R&D spillover parameter,  $\lambda$  as percentage of non-duplication and  $\nu$  as the degree of difficulty of R&D (see Dinopoulos and Segerström 1999). This slope coefficient can have any sign. Without duplication,  $\lambda = 1$ , and no difficulty,  $\nu = 0$ , we get the well-known case of Jones (1995) as  $g_y =$

$g_n/(1-\gamma)$ .<sup>6</sup> Segerström (1998) derives a similar result for quality improvements for consumer goods. Li (2000) gets this result with quality and variety improvements connected by cross-R&D spillovers. These equations have a zero intercept and without population growth there would be no income growth in the long run.

In the fully endogenous growth model of Howitt (1999) there is in addition a negative intercept. The steady-state growth rate in Howitt (1999) is  $g = \sigma\lambda n$ , with  $n$  as productivity-adjusted R&D expenditure,  $\sigma$  a spillover parameter and  $\lambda$  a Poisson arrival rate, all related to vertical R&D. From Fig.1 in Howitt (1999) one can solve for

$$n = \frac{(1-\alpha+\sigma)g_n/\psi(h^*)}{(1-\beta)(1+\frac{\alpha\sigma}{1-\alpha})} - \frac{r}{\lambda(1+\frac{\alpha\sigma}{1-\alpha})},$$

with  $\alpha$  as elasticity of production of intermediates,  $1-\beta$  as marginal cost of horizontal R&D expenditure diminished by subsidy  $\beta$ ,  $\psi$  as intensive-form production function of horizontal innovation,  $h^*$  the steady-state share of expenditure on research,  $r$  as the interest rate and  $g_n$  as population growth rate. Inserting the formula for  $n$  into that for  $g$  it follows that there is a negative intercept. Insertion of the standard consumption function  $g_c = (r-\rho)/\mu$ , with  $\rho$  as rate of time preference and  $\mu$  as consumption elasticity of marginal utility, after solving it for  $r$  and then solving for the steady-state value  $g = g_c = g_y$  again, does not change this qualitative result of a negative intercept (see Appendix). The generalised version of Segerström (2000) also requires positive population growth to have technical change and income growth.<sup>7</sup> Papers looking only at the R&D production function are unlikely to see the sign of the intercept as a crucial criterion to distinguish models because they ignore other parts of the model and their links with the R&D production function.

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<sup>6</sup> Note that for  $\gamma \leq 0$ , steady-state growth rates should be equal or even lower than population or employment growth rates, which seems to be unrealistic for high-income OECD countries. For 1977-2011 the employment in full-time equivalents grows at 1% per year; for 2001-11 this is less than 0.5%. Since 1995, OECD population growth rates are below 0.8%. Growth rates of GDP per capita are larger than these values, not smaller.

<sup>7</sup> This follows from equations (34) and (35) in Segerström (2000).

In an AK growth model, defined as having no decreasing marginal product of capital  $K$ , Dalgaard and Kreiner (2001) disaggregate  $K$  into a Cobb-Douglas function of two production factors, human capital  $H$  and productivity  $A$ , and assume a positive savings rate for the change of both. The result is a slope coefficient of  $g_n$  of minus unity (divided by the consumption elasticity of marginal utility) in case of a non-Benthamite utility function (no  $N_t$  multiplied to the per capita utility function) and a zero slope in the Benthamite case. In both cases, they have a positive intercept. A similar result is obtained by Young (1998) and Dinopoulos and Thompson (1998) with vertical spillovers and no horizontal spillovers (or neutralised ones). Peretto (1998), neutralising the horizontal knowledge spillover by decreasing market shares also belongs to this class of models. Peretto (2018) generalises some of these models in a way that fully and semi-endogenous growth theories are included as special cases. However, in these Schumpeterian fully endogenous growth models, unlike the Dalgaard/Kreiner result, there is a positive effect of the population growth rate on the growth rate of per-capita GDP or consumption as in the Lucas models. Smulders and van de Klundert (1995) find a zero slope for the case where neither research productivity nor knowledge spillovers depend on the number of varieties or the labour force.

Besides the standard representation of semi- versus fully endogenous growth models, our selected literature survey shows that Howitt (1999) differs through a negative intercept from the other fully endogenous growth models; duplication can destroy semi-endogenous growth in Dinopoulos and Segerström (1999); and the Bentham case of Dalgaard and Kreiner (2001) allows for zero effects of population growth as does the model of Smulders and van de Klundert (1995). This differentiation is important because it relates the papers to the question of expecting

positive or zero growth for the future in case that population growth goes to zero. Our paper therefore can be read as an empirical effort to figure out which of the special cases is more realistic.

#### TABLE 1 OVER HERE

Table 1 collects these cases of hypotheses for the relation between population growth and growth of income per capita in columns 1 to 3.<sup>8</sup> The last column shows the country-specific results from our empirical modelling approach, which we explain below. We want to indicate that in principle every result for the slope or the intercept alone is possible according to the theoretical models of Table 1, provided they allow for non-negative growth rates in line with the history of economic growth. The latter requirement can only be fulfilled if either the slope or the intercept are non-negative.

### 3 Estimation: Methods and results

#### 3.1 *Aspects of dynamic empirical modelling and the data*

When trying to test the hypotheses for slopes and intercepts of the theoretical models we must take into account that the assumption of exogenous population growth in large parts of endogenous growth theory may be a simplifying one. Human capital or income have an impact on population growth in rich and poor countries, which unified growth theory considers (Galor and Weil 1999). This should be taken into account when evaluating non-scale models<sup>9</sup>. For large panels, which include poorer countries than our panel, it is the log-level of income with several lags that

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<sup>8</sup> Tournemaine (2007), Bucci and Raurich (2017), Chang et al. (2017), and Bucci et al. (2018) survey the relevant literature more broadly. In a purely technical sense, the case of a zero slope and a positive intercept is also that of the neoclassical exogenous growth model. However, there is consensus that growth is generated by human capital and R&D and the question is just 'how exactly in detail'.

<sup>9</sup> The expression 'non-scale' emphasises the contrast with models that have the level of the population or human capital on the right-hand side of the formula for the solution of the growth rate as in Shell (1967) and others later. We do not consider this type of models. See Bloom et al. (2017) for a critique.

has an impact on population growth (Kelley and Schmidt 1995; Herzer et al 2012; Fosu et al 2016). We pose the question here for the growth rate solution of rich countries producing technical change, and therefore the (log-) level of income is not an aspect of the discussion. Instead, the structure of endogenous growth models is crucial and therefore the growth rates of population or labour and per-capita income, proportional to that of productivity in the long run, do matter. This implies the possibility of two-way causality between the growth rates of per capita income and the population,  $g_y$  and  $g_n$ .

Ha and Howitt (2007) use a simple error-correction model and favour the fully endogenous growth models based on an analysis of unit roots, cointegration and forecasting for the USA. They do not analyse the empirics of the intercept though and they do not estimate a dynamic model allowing for only endogenous variables using lags, both of which we will do here. Moreover, the relation between the two growth rates is not only a behavioural one, but rather it is based also on the labour constraint of the economic models. Therefore, we do not emphasise Granger causality but rather empirical relations with two-way causality in dynamic econometric models. Brander and Dowrick (1994) have pointed out that it is important to have a dynamic system analysis for this interaction to which we turn below.

We do not use TFP data because the process of making them tends to impose a Cobb-Douglas function for output production, which has a unit elasticity of substitution between capital and labour, where in general, many would prefer a lower one; however, there is no agreement to a certain value below unity. We consider the growth rates of per capita GDP (in constant 2005 US dollars) and population of 16 OECD countries for the period 1960-2014 taken from World Development Indicators.

Chang et al. (2017) use data for a longer period, but the volume and effect of R&D may be too small for some countries in earlier periods.

### 3.2 Pooled mean group estimation

Pooled mean group (PMG) estimators, assuming a common coefficient in the long-run relation for the two causality directions, estimated separately, result in long-term relations of  $g_y = -0.8g_n$  (Table 2, regression (1)) and  $g_N = 0.27g_y$  (Table 2, regression (2))<sup>10</sup>. They are very much different from each other and indicate two-way causality with opposite signs. We demean the variables in order to have time and period fixed effects in the long-term relation before running the regressions (1)-(3) and (5) of Table 2.<sup>11</sup> The first of these equations can be seen as an estimate of the equation used as framework by Ha and Howitt (2007) with TFP as endogenous variable. In addition, the first of these equations is very similar to the cross-section results in Strulik et al. (2013) using TFP growth rates. However, cross-section analysis does not take into account that in many countries growth rates of TFP and population are both falling during the after-war period. This latter fact suggests a positive correlation which we find in the second equation above and also if we use  $g_n(-20)$  in the first equation resulting in a long-run relation of  $g_y = 0.67g_n(-20)$  (not shown) with a large loss of observations though.<sup>12</sup>

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<sup>10</sup> These regressions should neither be mixed up with those regressing population growth or fertility on (several lags of) the log of GDP per capita (see Ahituv (2001), Herzer et al (2012) and Fosu et al (2016)) nor with growth regressions which have a lagged dependent variable (Li and Zhang 2007), or panel Granger causality studies of Chang et al. (2017). All of these studies address out-of-steady-state purposes, and growth regressions are mostly not made for endogenous growth models.

<sup>11</sup> The pooled mean-group estimator takes cross-section specific effects into account (Asteriou and Hall 2015). Modifying the variables by cross-section demeaning therefore does not change the results. However, demeaning for period fixed effects makes the first coefficient smaller and the second larger.

<sup>12</sup> Strulik et al (2013) in their Figure 2 show a positive correlation before 1920 and after 1970, and a negative correlation in between. The debates of modern growth theory focus on the phase since 1950 or 1960, because

TABLE 2 OVER HERE

3.3 Panel vector-error-correction models for two-way causality

In order to find a long-term relation with two-way causality, we use an econometric method that takes into account two-way causality. Cointegrated vector-autoregression (CVAR) or vector-error-correction (VEC) models are doing this.

VECMs have also the advantage that long-term relations may reflect steady-state results (Pesaran 1997, Breitung 2005) although economies are not yet in the steady state at least in the beginning of the estimation period. Lagged impacts are used to estimate the model, which is an aspect which Ha and Howitt (2007) regret to omit. We estimate three VECMs for our sample of 16 countries allowing for successively more heterogeneity.

In order to take fixed effects for countries and periods into account in the long-term relation of vector-error-correction model (1a, b), we subtract again the country and period specific averages of each variable from the variable itself and add the sample mean (see Greene 2012, p.404). The first VECM with four lags is as follows.

$$d(g^*_{yt}) = \alpha_1 [g_{yt-1} - \beta g_{nt-1} - \mu - c_i - \delta_t] + \gamma_{11}d(g^*_{yt-1}) + \gamma_{12}d(g^*_{nt-1}) + \gamma_{13}d(g^*_{yt-2}) + \gamma_{14}d(g^*_{nt-2}) + \gamma_{15}d(g^*_{yt-3}) + \gamma_{16}d(g^*_{nt-3}) + \gamma_{17}d(g^*_{yt-4}) + \gamma_{18}d(g^*_{nt-4}) + u_{1t}, t=1, \dots, T \quad (1a)$$

$$d(g^*_{nt}) = \alpha_2 [g_{yt-1} - \beta g_{nt-1} - \mu - c_i - \delta_t] + \gamma_{21}d(g^*_{yt-1}) + \gamma_{22}d(g^*_{nt-1}) + \gamma_{23}d(g^*_{yt-2}) + \gamma_{24}d(g^*_{nt-2}) + \gamma_{25}d(g^*_{yt-3}) + \gamma_{26}d(g^*_{nt-3}) + \gamma_{27}d(g^*_{yt-4}) + \gamma_{28}d(g^*_{nt-4}) + u_{2t}, t=1, \dots, T; \quad (1b)$$

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before this time education and invention were limited to a small part of the population. Ziesemer (2016) uses a cubic specification for labour force growth in a growth regression for a large panel of countries producing a negative effect only if labour growth is stronger than 2.45% as it is realistic for some African countries.

A ‘\*’ indicates that the variables are transformed in this way. All data are pooled in this model and the covariance matrix is  $E(u_t u'_t) = \begin{bmatrix} \sigma_{11} & \sigma_{12} \\ \sigma_{21} & \sigma_{22} \end{bmatrix}$  (see Canova and Ciccarelli 2013). This 2x2 covariance matrix means that no country interaction is taken into account because the data are pooled.

The second model version estimates jointly N=16 VECMs, one per country  $i$ :

$$d(g^*_y)_{it} = \alpha_{1i} [g_{y\ it-1} - \beta g_{n\ it-1} - \mu - c_i - \delta_i] + \gamma_{11i} d(g^*_{y\ it-1}) + \gamma_{12i} d(g^*_{n\ it-1}) + \gamma_{13i} d(g^*_{y\ it-2}) + \gamma_{14i} d(g^*_{n\ it-2}) + u_{1it}; \quad i=1, \dots, N; t=1, \dots, T; \quad (2a)$$

$$d(g^*_n)_{it} = \alpha_{2i} [g_{y\ it-1} - \beta g_{n\ it-1} - \mu - c_i - \delta_i] + \gamma_{21i} d(g^*_{y\ it-1}) + \gamma_{22i} d(g^*_{n\ it-1}) + \gamma_{23i} d(g^*_{y\ it-2}) + \gamma_{24i} d(g^*_{n\ it-2}) + u_{2it}; \quad i=1, \dots, N; t=1, \dots, T; \quad (2b)$$

Again, a ‘\*’ indicates double-demeaned variables, which are also used in the long-term relation in the implementation; period and country fixed effects can then be retrieved after the estimation (Greene 2012). All coefficients are country-specific with the exception of  $\beta$ , which is constrained to be identical over countries and periods as in the pooled mean-group (PMG) estimator. The covariance matrix then has format  $Nk \times Nk = 32 \times 32$ , as there are  $N = 16$  countries with each  $k=2$  variables and equations (see Groen and Kleibergen 2003, eq. (6)). A similar model of this type with common long-term coefficients for all countries is Breitung (2005). Our second model is a special case of this as we have at most one long-term relationship by assumption, confirmed by the Johansen-Fisher test below rejecting the hypothesis that the cointegration rank is  $r = k = 2$ . This approach can take into account interaction between the residuals of the countries when using the seemingly-unrelated-regression equations (SURE) and 3SLS estimators.

The third model differs from (2a, b) in that it leaves also the slope parameters of the long-term relation  $\beta_i$  free, except for being the same in the two equations of the VECM of each country, of course, and it has a country-specific constant in the long-term relation.

$$d(g_y)_{it} = \alpha_{1i} [g_{y\ it-1} - \beta_i g_{n\ it-1} - \mu_i] + \gamma_{11i} d(g_{y\ it-1}) + \gamma_{12i} d(g_{n\ it-1}) + \gamma_{13i} d(g_{y\ it-2}) + \gamma_{14i} d(g_{n\ it-2}) + u_{1it}; \quad i=1, \dots, N; \quad t=1, \dots, T; \quad (3a)$$

$$d(g_n)_{it} = \alpha_{2i} [g_{y\ it-1} - \beta_i g_{n\ it-1} - \mu_i] + \gamma_{21i} d(g_{y\ it-1}) + \gamma_{22i} d(g_{n\ it-1}) + \gamma_{23i} d(g_{y\ it-2}) + \gamma_{24i} d(g_{n\ it-2}) + u_{2it}; \quad i = 1, \dots, N; \quad t = 1, \dots, T; \quad (3b)$$

There are now country-specific slopes and intercepts in the long-term relation. The model has the country-specific long-term relations and the restriction of having no variables of other countries in the equations in common with the panel VEC model by Larsson and Lyhagen (2007). Whereas the latter impose the assumption of identical cointegration rank for all countries, our more traditional simultaneous equation model allows for having one or no cointegrating relation for each country. Maddala and Kim (1998, ch.5) discuss the relation and differences between the Johansen-Juselius (JJ) VECM approach, which we have used frequently elsewhere, and the traditional simultaneous equation approach. The latter has all dynamic information for the short and the long run when the cointegrating equations are specified. The JJ approach is needed mainly when the number and specification of cointegrating equations are not known and finding it adds valuable information. In our case, the maximum number of long-term relations is unity, given by theory, and allowed to be statistically insignificant, implying that we use the available information in the two variables. Chang et al. (2017) use Kónya's (2006) VAR version of the model (3a), (3b) for Granger causality analysis with full heterogeneity, or, conversely,

we use the VECM version of their VAR, with country-specific lag length in some versions of the model. Note that panel Granger causality tests also have an interpretation related to short and medium term business cycles, not only to steady-state growth rates and therefore we do not use it here but rather try to isolate steady-state growth relations.

In all three models, there is a constant but no time trend in the long-term relation and no constant outside the long-term relation as growth rates are assumed constant in the long run and their change then is zero.

### *3.4 Panel Unit roots and cointegration*

The variable for population growth has a common unit root according to the IPS and Fisher tests not taking into account cross-section dependence. For the cross-sectionally augmented CADF approach of equation (6) in Pesaran (2007), the average t-value is  $CIPS = -2.32$ , just allowing to reject the unit root hypothesis at the 5 percent level (Table II(b) in Pesaran 2007 shows the critical values). On the individual level, using Table I(b) in Pesaran (2007), we reject the individual unit root hypothesis only for three countries, AUS, DEU, PRT.<sup>13</sup> In line with this result, the chi-square Fisher test<sup>14</sup> rejects the hypothesis 'a unit root for all countries' at any standard level and beyond. With one exception, the coefficients for lagged population growth rates, which should be zero under a unit root, are negative, making near unit roots such as 0.95 instead of unity perhaps more likely. Overall, the population growth rates have a (near) unit root for most countries.

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<sup>13</sup> For additional lags we follow Baltagi (2013; ch.12), adding more lags of the differenced terms with common coefficients. PRT has a unit root if all lags are cross-section specific.

<sup>14</sup> The Fisher test statistic (called lambda-Pearson in Canning and Pedroni 2008), which is  $P = -2 \sum_1^N \ln p(i)$  (requiring using p-values at the highest precision in order to avoid logs of zero), follows a chi-square distribution.

The growth rate of the GDP per capita shows a unit root only for Italy and Portugal in the information on the alternative hypotheses on individual countries of the Fisher ADF test. The cross-sectionally augmented ADF test of Pesaran shows individual unit roots for FRA, DEU, NLD (and SWE if all lags are made cross-section specific).<sup>15</sup> The average t-value is CIPS = -4.055, rejecting the unit root hypothesis at any standard significance level. The chi-square Fisher test rejects the hypothesis that all countries have a unit root.

Income and population growth rates are cointegrated according to the standard panel cointegration tests of Kao for the null of a common unit root in the residuals. In the case of the Kao test, which takes into account country fixed effects, the probability of no cointegration or a common unit root in the residuals goes from 4.5 to 13.5 percent if the variables are time-demeaned.

All of the seven cointegration tests of Pedroni, which also take into account fixed effects and use country-specific autocorrelation coefficients, show cointegration; one test at the ten percent level, one at the two percent level and all others at the one percent level. The country-specific information on the heterogeneous alternative suggests no cointegration for DNK, ITA, SWE, UK, USA.

The Johansen-Fisher panel cointegration test based on the Fisher statistic from the combined trace and maximum eigenvalue tests rejects the hypothesis of having no cointegration and accepts the hypothesis 'at most 1 cointegrating equation', thus ruling out  $r = k = 2$  and estimation of VARs in levels. According to the individual tests, there is no cointegration for DNK, GRC, IRL, ITA, PRT, SWE, which is similar but not identical to the previous list. As the Johansen-Fisher test is based on separate

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<sup>15</sup> These results are in line with rejecting the null of a unit root too often when cross-section dependence is ignored (Gengenbach et al. 2009).

estimation of country-specific models, our joint estimation with the SURE method taking into account contemporaneous correlation may provide better results below.

Taking into account cross-section dependence in a panel cointegration test, we use the regression of  $g_y$  on  $g_n$  with country-specific coefficients in the correlated common effects (CCE) estimator of Pesaran (2006), which adds the period-specific averages across the countries for both variables (see equation (42) of Smith and Fuertes 2016). Running unit root tests on the residuals of this regression we find cointegration for all countries in the Fisher PP test, and unit roots for Netherlands and Portugal in the Fisher ADF and IPS tests (at the 1 percent level also for Finland and Ireland for both tests).<sup>16</sup> The presence of cointegration, when taking into account cross-section dependence, provides strong support for the use of a vector-error correction model.

Overall, the cointegration tests suggest that we should expect heterogeneity with mostly statistically significant relations between the two growth rates with a few exceptions where we found no cointegration in some of the tests. However, the standard tests have been developed for  $I(1)$  variables which we do not have throughout. What remains to be determined by our model then are the country-specific signs of slopes and intercepts and the average coefficients. We have to estimate VECM models (1a, b)-(3a, b) for several reasons. First, we have a mixture of stationary and (nearly) nonstationary variables for both growth rates and a way to estimate them jointly is to use ADF or VECM methods (Pesaran 1997; Maddala and Kim 1998, ch. 5; Pesaran 2015, p.550<sup>17</sup>) and check that residuals have no unit roots;

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<sup>16</sup> All coefficients that should be zero (one) under a unit root are lower, again making a near unit root hypothesis even likely as that of a unit root.

<sup>17</sup> "... the modelling procedure is robust to uncertainties surrounding the order of integration of particular variables. It is often difficult to establish the order of integration of particular variables using the techniques

second, we have cointegration for some countries and not for others. Some VAR-type of models, which are appealing for us because of the suggested two-way causality, have the disadvantage of assuming the same number of cointegrating equations for all countries (Groen and Kleibergen 2003; Larsson and Lyhagen 2007), whereas our preliminary cointegration analysis above would suggest that they are different, either one or none. Breitung (2005) assumes that there are the same cointegrating equations for all countries, which again cannot be the case here if we have cointegration for some countries but not for the others. Third, other models assume weak exogeneity (see Choi 2015), which is inadequate here because both growth rates may be endogenous in the first instance.

### 3.5 Estimation results for Panel VECMs

We treat cross-section dependence using the SURE method (Smith and Fuertes 2016; Kónya 2006), which is also included in 3SLS estimation. In case adjustment coefficients turn out to be zero for some countries, we have equations in first differences. If only one adjustment coefficient is zero, we have weak exogeneity as special case without imposing it for all countries.

The result obtained from maximum likelihood estimation of (1a) and (1b) - a VECM with four lags (because the underlying VAR has five lags) - is a slope of 0.62 and an intercept of 1.87. Country-specific deviations from the latter are in the range of (-0.87, 1.03), leading to a range of growth rates of  $g_y$  in the interval from (1.0, 2.9) for the special case  $g_n = 0$  (see Table 2, regression (3)), implying positive growth if

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and samples of data, which are available, and it would be problematic if the modelling procedure required all the variables in the model to be integrated of a particular order. However, the observations above indicate that, so long as the  $rx1$  cointegrating relations,  $\hat{\varepsilon}_t = \hat{\beta}'y_{t-1}$ , are stationary, the conditional VEC model, estimated and interpreted in the usual manner, will be valid even if it turns out that some or all of the variables in  $y_{t-1}$  are  $I(0)$  and not  $I(1)$  after all." (Pesaran 2015, ch.22, Cointegration analysis, p.550).

population growth vanishes hypothetically. Other estimation methods lead to similar results. To get an idea of the statistical significance of the intercept we use the VECM without fixed effects transformation of the data. The underlying VAR has three lags according to all criteria and the VECM yields the result  $g_y = 0.63g_N + 1.81$  with t-values of 1.95 for the slope and 6.44 for the intercept in Table 2, regression (4). The similar slope and estimate as with fixed effects indicates that the impact of fixed effects is limited although they could partly control for non-representative years or countries. In sum, slope and intercept are positive, statistically significant and economically of reasonable order of magnitude with and without fixed effects.<sup>18</sup> This preliminary result is in line with the first type of fully endogenous growth models in Table 1 and not with the others listed there. By implication, if population growth goes to zero we will still have a growth rate of 1.8% as an average across the sixteen countries of our panel. This result is therefore in line with the empirical literature discussed above, assuming slope homogeneity and favouring fully endogenous models.

The essence of estimating systems (2a, b) and (3a, b), when the long-term coefficients are (not) common to all countries, is

- (i) to estimate the equations for each country, and
- (ii) to take into account heterogeneity at least in the short term relations in (2a, b), and also in the long-term relations in (3a,b), and
- (iii) to take into account the relation between the residuals of the different countries (Groen and Kleibergen 2003).

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<sup>18</sup> Both VARs are stable. Non-linear spillover effects can in theory lead to u-shaped or hump-shaped effects of population growth rates (Diwarkar and Sorek 2017). Depending on the education policies and the value of the population growth rate itself there can be positive or negative impacts of exogenous population growth rates on income growth rates for constant policies in theoretical models (Prettner 2014).

Moreover, the lagged dependent variables may suffer from the Nickell bias when  $T$  is small and therefore we use lags as instruments for all variables in the first instance. We use first the three-stage least squares estimator, which combines instrumental variables and the SURE estimator dealing with contemporaneous correlation and cross-section dependence (Smith and Fuertes 2016) and later only the SURE estimation method without instruments when assuming that  $T$  is large enough.<sup>19</sup> The result for the estimation of (2a,b), a PMG estimator in a VECM, using again the double demeaned variables in Table 2, regression (5), is a slope of 1.29 and an overall constant of 1.4, with country fixed effects in the interval of (-1.45, 1.23). This implies for one country that in the special case of  $g_n = 0$  the growth rate is about zero, but positive for all others. Regression (5) in Table 2 would clearly favour the first type of the fully endogenous growth model again. However, the PMG estimator of the slope may suffer from heterogeneity bias. Jusélius et al (2014) suggest separate country-by-country estimation, which has the advantage of determining the adequate lag length per country, but the interaction of growth between the countries is not taken into account. This interaction may be stronger among the OECD countries of our sample than the developing countries they consider and therefore we try to consider it by way of estimating equations (3a, b).

Next, we estimate the 16 VECMs of (3a, b) jointly, with all coefficients flexible, using lag length two for all countries in regression (6) of Table 2, and taking into account contemporaneous correlation of the residuals using 3SLS again. This yields an average slope of 1.23. We can attribute the insignificance of the slope and intercept

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<sup>19</sup> Using GMM-HAC leads to a 'near-singular matrix' warning.

parameters either to heterogeneity or to having the case of non-duplication and degree of difficulty outweighing each other when setting coefficients to zero.<sup>20,21</sup>

Regression (7) in Table 2 uses country-specific lag lengths from the standard tests done country by country, after testing for stability. The average slope and intercept values are then 0.86 and 0.48, which are positive and between the results obtained so far. As in regression (6), these mean-group estimates are statistically insignificant.

The Nickell bias is only relevant under slope homogeneity. In case of slope heterogeneity, we are left with the standard Hurwicz bias from lagged dependent variables being pre-determined because  $y(-1)$  depends on  $u(-1)$ . As we have about 50 observations, we could assume that the bias is small and rely on the consistency of the least-squares estimator regarding the Hurwicz bias (Ramanathan 2001).

Because of the contemporaneous correlation across countries, we use the SURE estimation method in regressions (8) and (9) which otherwise correspond to (7) and (5). In both cases, the move from 3SLS to SURE changes the sign of the slope, but it remains statistically insignificant; instruments may have done more harm than good in the previous regressions. In contrast, the intercept and the adjustment coefficients become statistically significant.

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<sup>20</sup> According to Baltagi (2006) pooled estimation seems to be much more robust than allowing for heterogeneity.

<sup>21</sup> The slope coefficients of regressions (5) and (6) with 32x32 covariance matrix in Table 2 are twice as high as those with 2x2 covariance matrix in regression (3) and (4). Using the GMM method with heteroscedasticity and autocorrelation consistent coefficients and standard errors (HAC) or Maximum-likelihood leads to a 'near singular matrix' warning, indicating that the determinant of the inverted matrix in the estimator is close to zero. This warning is absent in the 3SLS estimates, but it may be the reason for the higher estimates. The significance of the slope parameter in regression (6) is low, indicating either that the heterogeneity is strong or we are in the curse of dimensionality because of a low number of observations. Adding common factors yields a 'near singular matrix' warning more often and so do time dummies for structural breaks, both without leading to other interesting results.

### *3.6 Linking panel estimation results to endogenous growth models*

Regressions (3)-(5) with slope homogeneity assumption and statistically significant results would favour the first type of fully endogenous growth models of Table 1, because we have a positive slope and intercept, both statistically significant. For regressions (6) and (7) the interpretation depends on the decision to look at the signs, favouring fully endogenous growth models again, or emphasising the insignificance. When setting slope and intercept to zero because of the insignificance, (6) and (7) would suggest a semi-endogenous growth model with non-duplication outweighed by the degree of difficulty (see Table 1, last column) predicting zero growth in the steady state. SURE regression (8) with statistically insignificant average slope and significantly positive average intercept supports the Bentham version of the AK model by Dalgaard and Kreiner (2001) and Smulders and van de Klundert (1995). Regression (9) with slope homogeneity as in (5) supports the non-Bentham version of the AK model of Dalgaard and Kreiner (2001) with significantly negative slope and positive intercept with unknown significance of implicit fixed effects.

Summing up the regression results, at best regression (6) and (7) when putting statistically insignificant coefficients to zero would support a special case of semi-endogenous growth models, but with the unrealistic expectation of zero growth. If we do not put them to zero, (6) and (7) support the first type of fully endogenous growth models. Regressions (8) and (9) support fully endogenous growth of the models with zero slope by Dalgaard and Kreiner (2001), and Smulders and van de Klundert (1995). Overall, this is much support for endogenous growth and little for semi-endogenous growth.

### *3.7 Some properties of the preferred regression*

We consider the SUR estimation with country specific lag length as the most plausible estimation method with  $T = 55$  data points for the period 1960-2010. In addition, instrumental variables may have damaging effects (Wooldridge 2002). Therefore, we prefer regression (8) and explore it a bit more in Table 3 following Canning and Pedroni (2008). In columns 1, 4, 7 and 10 we present the coefficients of intercept, slope and adjustment coefficients. Columns 2, 5, 8 and 11 show the t-values and 3, 6, 9 and 12 the p-values.

#### TABLE 3 OVER HERE

The intercepts are all positive except for the USA. By implication, all countries except for the USA would have an expected positive growth rate under zero population growth. However, five of the intercepts are not statistically significant. The average of the t-values found is 4.21 and thereby we have a statistically significant group mean of 2.5% in the absence of population growth according to the Fisher (lambda-Pearson) test. The t-values are so high that the chi-square distribution has a very small likelihood for the  $2N=32$  degrees of freedom and we reject the corresponding null hypothesis of a zero intercept.

The slopes are negative for nine countries, of which four insignificant, and positive for seven countries, of which three insignificant. The average t-value is almost zero and suggests a group mean of zero, as in regression (8) in Table 2 but without using the coefficient covariances in the calculation of the standard error. The lambda-Pearson/Fisher test statistic is strongly positive suggesting that the slopes are not always zero. The countries with insignificant slopes are similar to those found through the Johansen-Fisher panel cointegration tests above, with DEU now instead

of IRL there. Using contemporaneous correlation in the SUR method in addition to the separate VECMs in the Johansen-Fisher approach changes the results slightly.

The adjustment coefficients going to  $dg_y$  are all negative and statistically significant mostly at a lower level than one percent. The average t-value confirms this, and the lambda-Pearson test suggests that not all coefficients are zero. The adjustment coefficients going to  $dgn$  have positive or negative signs with a slightly positive average, with about half of them being statistically insignificant suggesting weak exogeneity of the population growth rate meaning that it does not react to disequilibrium in the long-term relations in half of all countries. The average t-value is just significant at the ten percent level and the lambda-Pearson statistic confirms that not all are zero. The average t-values and the lambda-Pearson are panel causality test statistics if there is cointegration under the null; the latter suggests that none of the four coefficients is zero in the whole panel.

### *3.8 Linking individual country results to endogenous growth models*

As the statistical insignificance of some coefficients of intercepts and slopes in regressions (3)-(9) may be due to heterogeneity, we want to look also at the country-specific results. In the last column of Table 1, we attribute country results to models besides related regressions, which show a similar result. In panel (a) we do this according to regression (7) using 3SLS, and in panel (b) according to regression (8) using SURE, both having country-specific lag length. In some cases an estimate is close to the ten-percent level, which we indicate by a note presenting the p-value. Nine to twelve countries have statistically significant positive or negative intercepts in the two panels of Table 1, which is in line with the panel-cointegration result reported

above and implies that they support fully endogenous growth theory. In contrast, five or six countries support semi-endogenous growth theory. Whereas the averaging regressions (3)-(5) favour the fully endogenous growth model, actually, according to the approach of this paper, no country falls into this class unless one goes to significance levels worse than ten percent, and only two or none fall into the class of a negative intercept. Among the fully endogenous growth cases, the scale neutral and non-Bentham model of Dalgaard-Kreiner (2001) dominate in the sense that the largest number of cases has a significantly positive intercept and a zero or negative slope. An implication of this is that growth can also remain positive if population growth goes to zero if intercepts are positive. In the case of a negative intercept, sufficiently strong population growth is required to get positive growth rates and has to outweigh the negative intercept constituted by interest costs or consumption growth. Finally, for one or four countries, we find slope and intercept statistically insignificant, which we interpret as semi-endogenous growth with non-duplication outweighed by the degree of difficulty. However, having Australia, Sweden and the USA in this class predicts zero long run growth for them if population growth vanishes. In contrast, Table 1, panel (b), predicts that these countries will have positive growth rates in the long run even if population growth goes to zero.

Our result provides also empirical support for Cozzi's (2017) hybrid model, which is based on a linear combination of the semi-endogenous and a fully endogenous growth model, leading to results of the latter when population growth is low as it is in our sample.

Overall, the averaging regressions (3)-(7) support the positive intercepts and slope coefficients of population growth in line with the fully endogenous growth model with or without statistical significance. However, under full heterogeneity in regressions

(7) and (8) with country-specific lag length, there are only two or even no country cases belonging to the class of the fully endogenous growth model with positive slope and intercept, but rather most countries have positive intercepts and zero or negative slopes as in the Dalgaard-Kreiner and Smulders-vdKlundert (1995) models. In contrast, regressions (3)-(5), imposing slope homogeneity with or without fixed effects, find statistically significant positive slopes and intercepts, but regression (9), the SURE version of (5) finds a significantly negative slope. When tests in our country panels are based on testing the slope and intercept of steady-state growth formulas for endogenous growth in VECMs, none of the growth models can be ruled out without abandoning the corresponding methods of handling heterogeneity and estimation. However, a negative intercept gets the least support. The preferred regression (8) supports statistically significant positive intercepts, and about equally many positive, negative or zero slopes. Overall, this supports fully endogenous growth models, but not necessarily those with a positive slope.

### *3.9 Econometric endogeneity*

The evaluation of regression (7) using 3SLS so far is based on sign and significance in the long-term relationship of economic theory only. The econometric definition of endogeneity would require that slope and adjustment coefficients are statistically significant. With this requirement, only income growth of the UK and population growth of Canada and Denmark would be endogenous in our estimate of (7). For all other countries the econometric conclusion within the 3SLS approach is that growth is weakly exogenous, when the emphasis focusses on the heterogeneity of regression (7). In contrast, in regression (8) using SURE, the income growth

equations have statistically significant adjustment coefficients throughout, meaning that income growth reacts to dis-equilibrium deviations from the long-term relation. Population growth is weakly exogenous for half of all countries (DNK, ESP, FRA, IRL, ITA, SWE, UK, USA) having statistically insignificant adjustment coefficients in the second equation for population growth dynamics. Therefore, it does not react to dis-equilibrium but only to changes in both growth rates for these countries whereas for the other eight countries it does react. Therefore, regression (8) fulfils the necessary condition for cointegration of having at least one non-zero adjustment coefficient per country (Canning and Pedroni 2008).<sup>22</sup>

The Johansen-Fisher panel cointegration test above showed no cointegration for DNK, GRC, IRL, ITA, PRT, SWE. Table 3 shows a zero slope for these countries except for IRL but instead also for DEU. This shows that taking cross-section dependence into account through the SURE method but not in the Johansen-Fisher test changes results slightly. In these cases with zero slope, our model suggests  $g_y(i) = c(i) > 0$  for five of these six countries (not GRC) with the same sign as in the fully endogenous growth models. Of the four countries with unit roots in the income growth variable, FRA and NLD have cointegration according to Table 3, but DEU and SWE have zero slopes. With only  $g_y$  and a constant in the long-term relation for DEU and SWE and strongly negative adjustment coefficients, this would imply that there is a near-unit-root variable  $g_y$  among the otherwise stationary variables or no unit-root variable in contrast to the CADF test of section 3.4.

Standard panel unit root tests (LLC, IPS, ADF-Fisher and PP-Fisher) for the 32 residuals of regression (8) show a probability of 0.0000 for the panel unit root

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<sup>22</sup> If an adjustment coefficient is zero, its product with the slope of the long-term relation is also zero, and then the variable plays no role in that equation.

hypotheses, indicating neither unit roots nor lack of cointegration and thereby that the regression is not spurious through the mixture of near unit roots and  $I(0)$  variables. We also use also Pesaran's (2007) unit-root test for the residuals correcting for cross-section dependence with no intercept and no trend. We find for the residuals of the income growth equations (3a) that the null of a unit root can be rejected for all countries but just not for Germany, with a coefficient of -0.47 instead of zero, and a t-value -2.14, where -2.28 is required for the 10% level (Pesaran 2007, Table Ia). Adding higher lags of differenced terms as suggested by Baltagi (2013, ch.12) does not change this. For the residuals of the population equations of (3b), a unit root cannot be rejected for the UK although the coefficient is -0.92 instead of zero. The average t-values, CIPS, of the panel unit root test are  $CIPS_1 = -6.64$  and  $CIPS_2 = -6.2$  for the two equations, rejecting the panel unit root hypothesis by far against the critical value of -1.86 at the one percent level (Pesaran 2007, Table IIa) and is support for our model. Pesaran's (2007) critical values are based on OLS, whereas our estimates reveal heteroscedasticity. If we use weighted least squares, our t-values go to a range of -3 to -9.3 for the residuals of the growth equations and -5.75 to -9.7 for the population growth equation making unit roots in the residual of our model even more unlikely. However, the Pesaran test is known to allow only for one common factor and its difference. Therefore, we use the fact that we have T is about 50 and N = 16, allowing to deal with cross-section dependence by use of the SURE estimation method (Smith and Fuertes 2016, Pesaran 2015), using the information about lag length that has been found before. For the residuals of the growth equations, we find t-values of -5.5 to -13.2 for the growth equation and -7.2 to -13.2 for the population growth equation. Both of these ranges are better than above and provide evidence against unit roots in the residuals of our model through mixing

$I(0)$  and near-unit-root variables.<sup>23</sup> Moreover, methods allowing for more common factor content in the residuals normally should bring up a higher likelihood for unit roots through cross-section dependence but in our case, this is not happening, indicating that cross-section dependence is weak and therefore should also not affect the estimate of the model (3a, b). Overall, we can conclude, that the residuals of our 16 VECMs have no unit roots and the combination of  $I(0)$  and (near) unit root variables is unproblematic (see footnote 16).

Canning and Pedroni (2008) emphasise in their appendix that  $-\alpha_{1i}/\alpha_{2i}$ , the negative of the ratio of the adjustment coefficients, is equal to the long-term effect of a transitory shock to population growth on income growth and its inverse for the opposite shock-effect relation under cointegration. The negative of the numerator, in column 7 of Table 3, is positive. The denominator, in column 10 of Table 3, is highly insignificant in six cases; the other ten cases have a positive value. However, in three of the ten cases (DEU, GRC, PRT) there is no cointegration because slope parameters in the long-term relation are highly insignificant. Therefore, for seven of the ten positive cases we have a positive long-term relation for the shock of population growth on income growth, and the other way around. When the coefficients in column 10 are essentially zero, the ratio is not well defined as we may not divide by zero, and the inverse is zero, implying having no impact of income growth on population growth in these six countries, and not in the three cases lacking cointegration. By implication, there is no impact in either direction in these cases. In short, a transitory shock to the residuals of changes in population growth rates has a long-run effect only in seven of sixteen countries. Heterogeneity is very strong here in both, slope coefficients and adjustment coefficients, although we are

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<sup>23</sup> We cannot apply spatial and multifactor methods for cross-section dependence here because they are valid for large  $N$ , a situation that we do not have here.

dealing only with rich countries. This result is in line with the mixed evidence for effects from population growth in general reported in the introduction.

### *3.10 Cross-unit cointegration*

A remaining econometric problem is the possibility of cross-unit cointegration among the income growth rates and among the population growth rates. Rivera-Batiz and Xi (1993) show that international capital movements make interest rates of countries similar to each other and thereby their growth rates. In other words, the interdependence may appear not only in the residuals with contemporaneous correlation as in the papers using Kónya's (2006) model, but also in the form of cointegration across countries. Although the correlation is spurious in the sense of depending on a third argument, interest rates, we cannot exclude the possibility that ignoring international interdependences biases our empirical results. Our estimation strategy is not a full-system analysis, defined as putting all variables of all countries into one VECM (Gonzalo and Granger 1995), but rather a combination of unit-by-unit analyses (Banerjee et al. 2004). It has the feature of a block-diagonal matrix of long-term coefficients (see Choi 2015) but it does not restrict the number of relations or the coefficients to be identical across countries. This literature is still under development (see Baltagi 2013; Pesaran 2015, Choi 2015). We stick here to the assumption of endogenous growth theory versions for closed and open economy models concerning the growth rates for income and population. Models with trade (Grossman and Helpman 1989) and international technology transfer (Howitt 2000) fit into the scheme of linearly related growth rate functions as well, where the two just

mentioned fit into the class of semi-endogenous models, no growth without population growth because R&D goes to zero and so do R&D spillovers.

#### TABLE 4 OVER HERE

Cross-unit cointegration in detail works as follows. In Table 4 we show the adjustment coefficients of the 16 VECMs in columns (1) and (2), where the highly insignificant ones have been set to zero. Dividing the values in columns (1) and (2) by those of column (1) yields columns (3) and (4). By implication column (3) consists of values 1 only, which in higher dimensional cases is the identity matrix, and column (4) consists of relative adjustment coefficients. Considering each country's row in columns (3) and (4) as a row vector, the country's row vector in columns (5) and (6) has been constructed such that the product of the row vector in (3), (4) and the transposed of the row vector in (5), (6) have product zero. The reader can check this by multiplying the element in column (3) with that in column (5), and those in (4) and (6) and adding up the two products, which gives zero for each country. The row vector in columns (5), (6) is called the orthogonal complement,  $\gamma'_{\perp}$ , of that of (3), (4) if the latter are transposed. Orthogonal complement to a vector  $\gamma$  is defined as  $\gamma'_{\perp}\gamma = 0$ , meaning that it has been constructed in a way that the product with  $\gamma$  is zero (see Lütkepohl 2005). Why do we need it?

Gonzalo and Granger (1995) consider the  $px1$  variables  $X$  - in our case,  $p=2$ , the growth rates  $g_y$  and  $g_n$  - of a VECM,  $X = \gamma\beta'X(-1) + \Gamma\Delta X(-i)$ , where  $\beta'X(-1)$  are the  $r$  cointegrating relations,  $\gamma$  ( $p \times r$ ) are the adjustment coefficients,  $\Delta X(-i)$  is a matrix of first differences of  $X$  with lags  $i$ , and  $\Gamma$  the matrix of its regression coefficients. They show that the data  $X$  can be decomposed into  $X = Af_t + B\beta'X$  with  $f_t = \gamma'_{\perp} X_t$ , ( $k \times p$ ,  $p \times 1$ , with  $k = p-r$ ), which is a real number per period, the series is  $I(1)$ , and  $f_t$  are not

*cointegrated*. In an example for six interest rates of Canada and the USA they first carry out the analysis as just described, leading to valid results as (i) the Johansen framework is used and (ii) specification errors are absent by assumption, and (iii) by construct the common components  $f_t$  are not cointegrated. Then they carry out the analysis separately for the three interest rates of Canada and the USA. They find the common components  $f_t$  for Canada and the USA and show that they are in fact cointegrated, hence they find 'cross-unit cointegration'. As they are not cointegrated when all six variables are analysed in one model rather than two, this means that separate analysis of twice three interest rates gives different results than joint analysis. As joint analysis with the Johansen-Juselius framework is free of objections the results of the separate analysis must be biased. Banerjee et al. (2004) suggest using this procedure as a bias test for cross-unit cointegration.

#### TABLE 5 OVER HERE

We have two variables,  $p=2$ , one cointegrating equation,  $r=1$ , and  $k = p - r = 2 - 1 = 1$ . We can multiply - for each country, as this corresponds to the separate analysis of the example above - the  $k \times p$  vector  $\gamma'_\perp$  with one row and two columns, namely (5) and (6) in Table 4, to each country's matrix  $X(t)$  with, in our case, two rows for the growth rates and one column. This leads to a linear combination of the growth rates for each country in each period,  $f_{it}$ . These common components are cointegrated for a VECM model with one lag, corresponding to two in the VAR, after dropping three countries with a low number of observations. The trace and maximum eigenvalue tests in Table 5 would suggest 9 or 4 cointegrating equations for the countries' common components at the 5 percent level as shown in Table 5, or, with one lag less and a time trend included, we would have 5 cointegrating equations at the six percent level for both tests (not shown). This would suggest that our results are

biased, as common components from separate models should not be cointegrated according to Gonzalo and Granger (1995). In order to check whether the bias is strong, one would need to have the joint analysis. Unfortunately, working with 2x16 variables is not possible in the VECM and therefore we cannot check the properties of the bias, but rather just admit its presence and hope that it is weak or that removing it would even strengthen our results. However, we would like to point out that analysis of cross-unit cointegration has not been done often in economics and is new in the economics of growth. It suggests that international interconnectedness should be taken into account as a next step in a later paper.

#### **4 Conclusion**

We have classified fully and semi-endogenous growth models in six categories according to the slopes and intercepts of the linear relation between income and population growth. We have tested endogenous growth models from the perspective of this relation. The crucial step for the empirical work is to give up the simplifying assumption of growth models that population growth is exogenous and test it instead. When population growth is endogenous and we allow for two-way causality, under the assumption of slope homogeneity a vector-error-correction approach shows that the relation between growth rates of income per capita and the population is positive with positive intercept. Both coefficients as well as the adjustment coefficients are statistically significant. Under the assumption of slope heterogeneity, adjustment coefficients of the ECM-term in the income-growth equation are always statistically significant. For Greece, slope and intercept are statistically insignificant. Intercepts are significantly positive when the slopes are zero

or insignificant in five cases, and the slopes are positive when the intercepts are zero or insignificant requiring positive population growth in the long-run in four cases. By implication, all countries fall into one of the categories of fully or semi-endogenous growth and results are therefore meaningful. However, under slope homogeneity they fall into the category with positive slope, but with heterogeneity they fall into the category of zero or negative slope. There are eleven countries with statistically significant positive intercepts, four with positive but insignificant intercepts and only one with negative, insignificant intercept. Therefore, both assumptions, slope homogeneity and heterogeneity, lead us to favour fully over semi-endogenous growth. We cannot exclude the cases where population growth has to be positive to continue growing. In particular, the results for the USA are in line with semi-endogenous growth models. Intercepts are predominantly positive and suggest that we can have endogenous growth also when population growth stops.

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**Table 1 Cases of long-run growth formulas**

**(a) Models and country classification from regression (7)**

<u>Hypothesis</u>	<u>slope</u>	<u>intercept</u>	<u>Country classification from regression (7)</u>
Fully endogenous (a)	positive	positive	averages from regressions (3)-(5)
Semi-endogenous $\lambda=v$ (b)	positive	zero	CAN? (c), FRA,
	zero	zero	AUS, GRC, SWE, USA; average regr. (6), (7)
Howitt 1999	positive	negative	CAN? (c), NLD
AK non Bentham	negative	positive	DNK, FIN, LUX, UK; ave. regression (9)
AK Bentham (d)	zero	positive	ESP, DEU, IRL, ITA, PRT; ave. regression (8)

(a) Includes Lucas 1988, Young 1998, Peretto 1998, Dinopoulos/Thompson 1998, Peretto 2018. (b) Non-duplication parameter equals degree of difficulty. (c) Intercept for CAN has  $p=0.13$ . (d) The special case in Smulders and van de Klundert (1995) should be most plausibly in this class but could also have a zero intercept.

**(b) Models and country classification from regression (8)**

<u>Hypothesis</u>	<u>slope</u>	<u>intercept</u>	<u>Country classification from regression (8)</u>
Fully endogenous (a)	positive	positive	AUS (c); FRA? (c); ave. regressions (3)-(5)
Semi-endogenous $\lambda=v$ (b)	positive	zero	NLD, FRA? (c), CAN, USA;
	zero	zero	GRC average regressions (6), (7)
Howitt 1999	positive	negative	-
AK non Bentham	negative	positive	UK, LUX, IRL, FIN, ESP; ave. regression (9)
AK Bentham (d)	zero	positive	AUS (c), DNK, DEU, ITA, PRT, SWE; ave. (8)

(a) Includes Lucas (1988), Young 1998, Peretto 1998, Dinopoulos/Thompson 1998, Peretto 2018.

(b) Non-duplication parameter equals degree of difficulty.

(c) AUS has a positive slope with  $p$ -value 0.12; for FRA intercept has  $p=0.18$ .

(d) The special case of high duplication in Smulders and van de Klundert (1995) should be most plausibly in this class but could also have a zero intercept.

Table 2 Error-correction estimates for growth rates of income and population										
Regression	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	
Dependent variable(s)	d(gy)	d(gn)	d(gy), d(gn)	d(gy), d(gn)	d(gy) <sub>i</sub> , d(gn) <sub>i</sub>					
<i>long run</i>										
dependent variable	gy	gn	gy	gy	gy	gy	gy	gy	gy	
regressor	gn	-0.8 (-3.46)	- (1.9)	0.62 (1.95)	0.63 (3.255)	1.29 (0.64) (k)	1.23 (i) (0.48) (k)	0.86 (i) (-0.73) (k)	-0.21 (i) (-0.73) (k)	-0.9 (5.47)
	gy	-	0.275 (3.5)	-	-	-	-	-	-	-
	constant	-	-	1.87+(-0.87,1.03)	1.8 (-6.45)	1.4+(-1.45, 1.23)	0.195 (g) (0.157) (k)	0.48 (g) (0.26) (k)	2.52 (g) (8.59) (k)	2.2+(-0.8, 1.33)
<i>short run</i>										
adjustment coeff.		-0.72 (-17.5)	-0.086 (-4.68)	-0.47, 0.0675 (-8.28), (7.7)	-0.42, 0.035 (-10.2), (7.28)	-0.39, 0.06 (f) (-2.227), (2.99) (k)	0.14, -0.018 (h) (0.12),(-0.157) (k)	-0.534, 0.035 (h) (-1.27),(0.06) (k)	-0.68, 0.033 (h) (-17.1),(7.9) (k)	-0.67, 0.03 (f) (-16.7), (6.4) (k)
regressors										
Lags (coeff. or no)	D(gn)	0.87 (-1.61)	2	4	2	2	2	<b>country specific</b>	country specific 2	
Lags (coeff. or no)	D(gy)	-	1	4	2	2	2	<b>country specific</b>	country specific 2	
Countries		16	16	16	16	16	16	16	16	
Period		1962-2014	1963-2014	1966-2014	1964-2014	1965-2014	1965-2014	1963-2014	1962-2014	1964-2014
Observation		828	814	764	796	1560	1560	1560	1592	1592
Method		PMG	PMG	<b>FIML</b>	FIML	<b>3SLS</b>	3SLS	3SLS	<b>SUR of (7)</b>	<b>SUR of (5)</b>
		ARDL(1,1)	ARDL(3,1)	<b>VECM</b>	VECM	<b>VECM, PMG</b>	VECM, <b>MG</b>	VECM, MG	VECM, MG	VECM, PMG
Log likelihood		-1561.246	347.6567	-1526.6, -97.38	-1833.9, -136.6	-	-	-	-	-
Fixed effects		countries	countries	countries	<b>none (d)</b>	countries	-	-	-	countries
		periods (a)	periods (b)	periods (c)		periods (e)	- (l)	- (l)	- (l)	periods (e)
p (adj Q) (j)		-	-	0.16; 0.55	0.04, 0.39	0.16, 0.68	0.066, 0.32	0.14, 0.62	0.16, 0.39	0.15, 0.61

t-values in parentheses.	Bold symbols indicate new elements compared to previous regressions in the table.
(a) Fixed effects via demeaning; adjustment coefficients range between (-1.02, -0.36; highest p=0.0001); coefficients for d(gn) range from -2.1 to 4.77. AIC for lag selection.	
(b) Fixed effects via demeaning; AIC for lag selection. Adjustment coefficients range between -0.26 and 0.088 with highest p-value =0.0019	
(c) Fixed effects via demeaning; covariances are $\sigma(1,1) = 3.185$ ; $\sigma(2,2) = 0.0755$ ; $\sigma(1,2) = \sigma(2,1) = -0.24$ . Lag length choice according to lag length criteria, stability test, and LM and Portmanteau serial correlation tests. Intercept of 1.868 plus country-specific deviations in the range of (-0.87, 1.03)	
(d) Lag length of VAR is three for all criteria. Boswijk (1995) for s.e. of long run coeff. . Covariances are $\sigma(1,1) = 5.87$ ; $\sigma(2,2) = 0.0825$ ; $\sigma(1,2) = \sigma(2,1) = -0.055$ .	
(e) Country and period fixed effects via demeaning; lag length fixed at 2; long term coefficient constrained to be identical for all countries; adjustment coefficients averaged over equations of 16 countries; other coefficients unconstrained. Covariance matrix is 32x32. Portmanteau test for null of no serial correlation has p-values for adj Q-stat (see Lütkepohl 2005) between 0.16 and 0.68.	
(f) average over first and second set of equations (2a), (2b).	
(g) average across countries with seven significantly positive, three significantly negative and six insignificant results for the intercept.	
(h) average across countries for first and second set of equations.	
(i) average over values for 16 countries.	
(j) lowest and highest p-value for 12 lags in the Portmanteau test for 'no serial correlation' hypothesis (see Lütkepohl 2005).	
(k) the variance of a mean group estimator is obtained as the sum over all elements in the coefficient covariance matrix divided by 256. Taking the square root yields the standard error that can be used to divide the coefficient to get the t-value. See Davidson and MacKinnon (2004), formula (3.68).	
(l) Under full heterogeneity time fixed effects are in the residuals.	

	1	2	3	4	5	6	7	8	9	10	11	12
<i>Country</i>	<i>Interc</i>	<i>t-value</i>	<i>p-value</i>	<i>slope</i>	<i>t-value</i>	<i>p-value</i>	<i>adj c dgy</i>	<i>t-value</i>	<i>p-value</i>	<i>adj c dgn</i>	<i>t-value</i>	<i>p-value</i>
AUS	1.34	2.25	0.025	0.58	1.545	0.123	-0.83	-4.76	0.0000	0.175	5.111	0.000
CAN	<b>0.86</b>	<b>0.99</b>	<b>0.323</b>	1.14	1.784	0.075	-0.76	-8.76	0.0000	<b>0.028</b>	3.652	0.000
DNK	2.23	3.03	0.003	<b>-0.60</b>	<b>-0.387</b>	<b>0.698</b>	-0.56	-4.11	0.0000	<b>-0.001</b>	<b>-0.190</b>	<b>0.849</b>
ESP	4.76	4.47	0.000	-3.29	-2.770	0.006	-0.28	-4.03	0.0001	<b>-0.003</b>	<b>-0.483</b>	<b>0.629</b>
FIN	4.65	4.85	0.000	-4.45	-1.901	0.058	-0.72	-5.64	0.0000	0.010	1.897	0.058
FRA	<b>0.72</b>	<b>1.32</b>	<b>0.187</b>	2.61	3.978	0.000	-0.53	-7.72	0.0000	0.011	1.441	0.150
DEU	2.06	13.18	0.000	<b>-0.26</b>	<b>-0.641</b>	<b>0.522</b>	-1.26	-7.23	0.0000	0.199	4.111	0.000
GRC	<b>1.42</b>	<b>0.94</b>	<b>0.348</b>	2.66	1.149	0.251	-0.24	-2.49	0.0129	0.021	3.558	0.000
IRL	7.76	5.38	0.000	-3.49	-2.923	0.004	-0.45	-2.49	0.0010	<b>-0.010</b>	<b>-0.980</b>	<b>0.327</b>
ITA	1.84	2.73	0.007	<b>0.27</b>	<b>0.207</b>	<b>0.836</b>	-0.46	-4.76	0.0000	<b>-0.001</b>	<b>-0.090</b>	<b>0.928</b>
LUX	4.26	7.53	0.000	-1.34	-2.995	0.003	-0.86	-7.19	0.0000	0.023	3.241	0.001
NLD	<b>0.57</b>	<b>1.02</b>	<b>0.309</b>	2.73	4.150	0.000	-0.50	-3.99	0.0001	0.015	2.454	0.014
PRT	2.68	4.17	0.000	<b>-0.01</b>	<b>-0.015</b>	<b>0.988</b>	-0.49	-4.85	0.0000	0.054	3.703	0.000
SWE	2.30	5.93	0.000	<b>-0.54</b>	<b>-0.815</b>	<b>0.415</b>	-0.78	-6.79	0.0000	<b>0.004</b>	<b>0.528</b>	<b>0.598</b>
GBR	2.89	9.74	0.000	-1.42	-2.418	0.016	-0.99	-10.34	0.0000	<b>-0.001</b>	<b>-0.303</b>	<b>0.762</b>
USA	<b>-0.09</b>	<b>-0.12</b>	<b>0.907</b>	2.10	2.859	0.004	-1.18	-9.36	0.0000	0.010	1.455	0.146
Average	2.52	4.21		-0.21	0.050		-0.68	-5.91		0.033	1.819	
lambda-Pearson		508.0	0.000		113.5	0.000		695.81	0.0000		139.2	0.000

Last p-values in a column belong to the lambda-Pearson/Fisher test statistic.  
Lambda-Pearson test statistic (below the t-values) is  $-2(\sum \text{from } l=1 \text{ to } l=16 \text{ over } \ln(p_l))$ .  
P-values in the formula for the lambda-Pearson test statistic are taken with highest possible precision to avoid logs of zeros.  
P- and t-values with strong statistical insignificance are printed in bold.

	(1)	(2)	(3)	(4)	(5)	(6)
	adj c dgy	adj c dgn	adj c dgy	adj c dgn	adj c dgy	adj c dgn
AUS	-0.835	0.175	1.000	-0.209	0.209	1
CAN	-0.756	0.028	1.000	-0.037	0.037	1
DNK	-0.562	0.000	1.000	0.000	0.000	1
ESP	-0.285	0.000	1.000	0.000	0.000	1
FIN	-0.723	0.010	1.000	-0.014	0.014	1
FRA	-0.528	0.011	1.000	-0.021	0.021	1
DEU	-1.259	0.199	1.000	-0.158	0.158	1
GRC	-0.243	0.021	1.000	-0.087	0.087	1
IRL	-0.454	0.000	1.000	0.000	0.000	1
ITA	-0.460	0.000	1.000	0.000	0.000	1
LUX	-0.862	0.023	1.000	-0.026	0.026	1
NLD	-0.499	0.015	1.000	-0.031	0.031	1
PRT	-0.487	0.054	1.000	-0.110	0.110	1
SWE	-0.780	0.000	1.000	0.000	0.000	1
GBR	-0.995	0.000	1.000	0.000	0.000	1
USA	-1.181	0.010	1.000	-0.008	0.008	1

(a) insignificant coefficients set to zero

**Table 5 Cointegration tests for potential common I(1) components from country models**

Sample (adjusted): 1963 2014,  
 Included observations: 52 after adjustments  
 Trend assumption: No deterministic trend (restricted constant)  
 Series: FCAN FDNK FESP FFIN FFRA FGBR FGRC FIRL FITA FLUX FNLD FSWE  
 FUSA  
 Lags interval (in first differences): 1 to 1

Unrestricted Cointegration Rank Test (Trace)

Hypothesised No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None	0.951958	689.6589	NA	NA
At most 1 *	0.869430	531.8032	348.9784	0.0000
At most 2 *	0.814283	425.9390	298.1594	0.0000
At most 3 *	0.784104	338.3954	251.2650	0.0000
At most 4 *	0.616982	258.6816	208.4374	0.0000
At most 5 *	0.589815	208.7787	169.5991	0.0001
At most 6 *	0.575511	162.4391	134.6780	0.0004
At most 7 *	0.512055	117.8819	103.8473	0.0043
At most 8 *	0.436031	80.56921	76.97277	0.0259
At most 9	0.375659	50.78593	54.07904	0.0953
At most 10	0.235514	26.29091	35.19275	0.3259
At most 11	0.135295	12.32625	20.26184	0.4200
At most 12	0.087599	4.767156	9.164546	0.3099

\*\*MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesised No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None	0.951958	157.8557	NA	NA
At most 1 *	0.869430	105.8642	77.38180	0.0000
At most 2 *	0.814283	87.54368	71.33542	0.0008
At most 3 *	0.784104	79.71372	65.30016	0.0013
At most 4	0.616982	49.90296	59.24000	0.3046
At most 5	0.589815	46.33958	53.18784	0.2113
At most 6	0.575511	44.55719	47.07897	0.0908
At most 7	0.512055	37.31270	40.95680	0.1215
At most 8	0.436031	29.78328	34.80587	0.1761
At most 9	0.375659	24.49502	28.58808	0.1529
At most 10	0.235514	13.96466	22.29962	0.4652
At most 11	0.135295	7.559092	15.89210	0.6010
At most 12	0.087599	4.767156	9.164546	0.3099

\*\*MacKinnon-Haug-Michelis (1999) p-values

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