International Migration with Heterogeneous Agents: Theory and Evidence

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Abstract

Temporary migration, though empirically relevant, is often ignored in formal models. This paper proposes a migration model with heterogeneous agents and persistent cross country income differentials that features temporary migration. In equilibrium there exists a positive relation between the stock of migrants and the income differential, while the net migration flow becomes zero. Consequently, existing empirical migration models, estimating net migration flows, instead of stocks, may be misspecified. This suspicion appears to be confirmed by our investigation of the cointegration relationships of German migration stocks and flows since 1967. We find that (i) panel-unit root tests reject the hypothesis that migration flows and the explanatory variables are integrated of the same order, while migration stocks and the explanatory variables are all I(1) variables, and (ii) the hypothesis of cointegration cannot be rejected for the stock model.

Keywords: International migration, temporary migration, panel cointegration. *JEL code*: C23, C53, F22.

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1 Introduction

Today's international migration – opposed to that of the 19th century – consists predominantly of temporary migrants, where an individuals' stay abroad may vary from a few months to several decades. Return migration in Europe makes up almost 10 per cent of migration stocks p.a.,[1] and micro studies indicate that around 80 per cent of the migrants in Europe eventually return to their home countries (Dustmann, 1995; Karras and Chiswick, 1999). A central implication of migration being temporary is that each migration stream is associated with a compensating counter-stream, as Ravenstein (1889) already observed in his famous 'Laws of Migration' at the end of the 19^{th} century.

Although often quoted, many of Ravenstein's fundamental observations have not been addressed systematically in the migration literature. Most traditional migration models treat migration as permanent. Starting with the seminal contributions of Hill (1987) and Djajic and Milbourne (1988) a number of models have meanwhile analysed the phenomenon of temporary migration, but these models usually ignore the heterogeneity of individuals. Consequently, even though migration is temporary, the same length of migration episodes applies for all agents.[2]

[1] See the evidence provided by SOPEMI (2003) for a number of OECD countries; for Germany see Bundesamt (2003).

[2] Hill (1987) and Djajic and Milbourne (1988) treat migration as an intertemporal optimisation problem, where the length of migration is endogenously determined by host and home wages and differences in utility between consumption abroad and at home. Yet as they employ the concept of a representative agent all migration decisions (including the length of stay) are identical for all agents. Building on these models, Dustmann and Kirchkamp (2002) and Mesnard (2004) consider problems such as liquidity constraints, differences in purchasing power parities across countries, and enhanced options for self-employment for return migrants. Moreover, Dustmann (1995), Dustmann and Kirchkamp (2002), Dustmann (2003) and Mesnard (2004) find micro evidence that the length of migration episodes depends on individual human capital characteristics. However, the observed heterogeneity across individuals is not considered by the underlying migration

This paper takes another route by considering heterogeneous preferences of individuals with regard to the choice of location. The basic set-up of the model is related to the standard model of temporary migration originally developed by Djajic and Milbourne (1988), but departs in several aspects: First, and most importantly, it is assumed that individuals discount consumption in foreign countries by a certain factor, which varies across individuals.[3] Second, in order to arrive at analytical solutions for the case of heterogenous agents with individual – and model endogenously determined – length of migration spells, we have to employ specific functional forms. Finally, the present model departs from Djajic and Milbourne (1988) with respect to wage reactions to migration flows. Arguably, the traditional assumption where wage equalisation determines the amount of migration in equilibrium appears to run counter to empirical evidence. Actual migration is small relative to persistent cross-country income differentials. Furthermore, institutional settings in Europe do limit wage reactions in both the receiving and sending countries. Nevertheless migration stocks are relatively stable over time despite these large wage arbitrage opportunities, this is for example the case within the EU and the European Economic Area (EEA), where migration barriers have been largely abolished. Based on these observations, equilibrium in the present model is not driven by the wage equating forces of migration, but instead by the heterogeneity of agents, which determines both the amount and duration of migration in reaction to a given income differential.[4]

models. One exception in the literature is the model of Stark (1995), which explains differences in migration duration by asymmetric information concerning the human capital characteristics of high and low productivity migrants. After the true type is revealed, low-productivity migrants are dismissed and return home, accordingly displaying a shorter migration duration than high productivity types.

[3]Some other models in the literature consider heterogeneous preferences as well (Faini and Venturini, 1995), but to the best of our knowledge the consequences of heterogeneous preferences have not yet been analysed in the context of temporary migration.

[4]Alternatively, the present analysis may be viewed as a partial equilibrium version of

Based on these assumptions, the model succeeds in distinguishing among three types of individuals within the population of the sending country: stayers, i.e. those who stay at home and do not migrate; temporary migrants, i.e. those who return home within their lifetime; and permanent migrants, i.e. those who migrate for their entire lifetime. Furthermore, within the group of temporary migrants, the duration of the migration episode varies across individuals. Accordingly migration flows and stocks are composed of a variety of different behaviours by agents. The average duration of migration episodes as well as the number of permanent migrants tend to increase with an increasing income differential. The stock of migrants, i.e. the share of the population which tends to stay abroad at a certain point of time, increases with the income differential between the host and the home country as well. Moreover, the net migration flow is zero in equilibrium. Gross migration flows remain, however, a positive function of the income differential.

These results have important implications for the estimation of macro migration models: The standard macro models of migration, following the famous Harris and Todaro (1970) model, presume that an equilibrium between net migration *flows* and the explanatory variables emerges. In contrast our model implies that an equilibrium relationship between migration *stocks* and the explanatory variables arises in the long run, while net floes become zero. Accordingly in the empirical part of the paper we test the hypothesis whether migration stocks or flows and the explanatory variables form a long-run equilibrium, or, in technical terms, a cointegration relationship, empirically. Our analysis is based on migration to Germany from EU source countries during the period from 1973-2001. Note that the EU is a natural laboratory for studying international migration behaviour, since institutional barriers for

more traditional set-ups.

migration have been removed there since the late 1960s. Following the Engle and Granger (1987) procedure, we test first whether the variables of the stock or the flow model form a cointegrated set. For this purpose, we apply panel unit-root and panel cointegration tests, which increase the statistical power in comparision to univariate unit-root and cointgration tests. We find that migration flows are stationary variables while the explanatory variables such as income and employment variables are integrated of the first order (I(1) variables). Thus the hypothesis of the traditional migration model in the empirical literature that migration flows and the explanatory variables form a cointegrated set, is not supported by our data set. In contrast, we find that migration stocks are I(1) as well. Moreover, our panel cointegration tests suggest that the hypothesis that migration stocks and the explanatory variables form a cointegrated set cannot be rejected.

The remainder of this paper is as follows. Section 2 presents the formal model of migration with heterogeneous agents. It derives the amount of permanent and temporary migration as well as the individual and aggregate duration of migration episodes and presents results for the implied migration stocks and flows. In Section 3 we first discuss the alternative flow and stock specifications for empirical macro migration models, and then apply panel unit-root and panel cointegration tests in order to prove whether the variables of the alternative models form a cointegrated set. In section 4 the stock model is estimated, in particular the cointegrating vectors and the short-run dynamics of the stock model are estimated by employing an error correction model. Section 5 concludes.

2 A migration model with heterogeneous preferences

2.1 The model

Consider an economy where at each instant in time, t, there are N individuals i born, endowed with one unit of labour each, and who each live for the same period of time, T_i, normalised to 1, i.e. $T_i = 1, \ \forall \ i = 1, ..., N$. Each individual is continuously employed throughout his or her life but has the choice of staying abroad for a period τ_i, where $0 \leq \tau_i \leq 1 \ \forall \ i = 1, ..., N$. As in Djajic and Milbourne (1988), agents make and execute their migration decision at time $t_i = 0$. In the foreign country, each domestic and migrant worker receives the income level y^*, in the sending country the income level is y, where $y < y^*$, i.e. outward migration only occurs from home to foreign country.[5] The utility flows which individuals perceive from consumption (that is, living) at home and abroad respectively are given by:

$$
\begin{aligned}
u(c_i) &= c_i^{\alpha} && (1) \\
u^*(c_i^*) &= \gamma_i^{1-\alpha} c_i^{*\alpha}, && (2)
\end{aligned}
$$

where c_i and c_i^* are consumption at home and abroad, respectively, α $(0 < \alpha < 1)$ is a parameter of the utility function, identical for all agents, and $\gamma_i \in [0, 1]$ is a preference parameter, which is heterogenous across agents. The parameter γ_i captures the fact that individuals receive less utility from consumption abroad than at home. The utility functions in (1) and (2) display the feature that the marginal utility enjoyed from the same rate of consumption is higher at home than abroad, i.e. that $u^{*\prime}(x) < u'(x)$, thus

[5] Variables with an asterisk denote throughout the Section values in the foreign country.

fulfill the conditions laid out in Djajic and Milbourne (1988).[6]

The lifetime utility of a migrating individual returning to the home country at time τ_i can then be written as[7]

$$V_i = \tau_i \gamma_i^{1-\alpha} c^*(t)^\alpha + (1-\tau_i) c(t)^\alpha. \tag{3}$$

The intertemporal maximisation problem of the individual is then straightforward (see e.g. Djajic and Milbourne (1988)): choose the duration of the stay in the foreign country, τ_i, and the rates of consumption over time abroad, $c_i^*(t)$, and at home, $c_i(t)$, such that lifetime utility (3) is maximised subject to the budget constraint

$$\tau_i y^* + (1-\tau_i) y - \tau_i c_i^*(t) - (1-\tau_i) c_i(t) \geq 0. \tag{4}$$

The first-order conditions (see Appendix A.1) give rise to the following relations:

$$\begin{aligned}
\alpha \gamma_i^{1-\alpha} c_i^*(t)^{-(1-\alpha)} &= \lambda, && (5)\\
\alpha c_i(t)^{-(1-\alpha)} &= \lambda, && (6)\\
\gamma_i^{1-\alpha} c_i^*(t)^\alpha + c_i(t)^\alpha &= \lambda \left(c_i^*(t) - c_i(t) + y - y^* \right), && (7)\\
\tau_i \left(y^* - y + c_i(t) - c_i^*(t) \right) &= c_i(t) - y. && (8)
\end{aligned}$$

Since the shadow value of wealth, λ, is time-invariant, (5) and (6) imply that

[6] As is usual in the literature, we interpret the condition that $u^{*\prime}(x) < u'(x)$ to capture the fact that closer social relations to friends and relatives in the home country, a familiar cultural environment and other factors associated with the home country result in a higher utility for the same rate of consumption in the home country (Faini and Venturini, 1995).

[7] Notice that we ignore discounting by setting the discount factor implicitly to one. Nevertheless, none of the results below depend on this assumption, see e.g. Dustmann (2003), Dustmann and Kirchkamp (2002), and Mesnard (2004) for a similar approach. Discounting is, however, included in Djajic and Milbourne (1988).

$c_i^*(t) = c_i^*$ and $c_i(t) = c_i$, $\forall\, i = 1, ..., N$, i.e. consumption at home and abroad is constant over time. Moreover, equating the left-hand side of (5) and (6) gives:

$$c_i^* = \gamma_i c_i \,. \tag{9}$$

Thus consumption during the migrants stay abroad is a fraction of the consumption upon his/her return to the home country. Next, from (9) and (7), after substituting λ from (6), we are able to solve for c_i:

$$c_i = \frac{\alpha}{1-\alpha}\frac{1}{1-\gamma_i}\left(y^* - y\right) \,, \tag{10}$$

i.e. consumption at home – and thus via (9) also consumption abroad – is a linear function of the income differential.[8]

Finally, using (9), (10) and (8) one can calculate the optimal length of a migrants's stay abroad:

$$\tau_i = \frac{\alpha}{1-\gamma_i} - \frac{(1-\alpha)y}{y^* - y}. \tag{11}$$

This optimal duration of migration displays the following reactions to changes in the various parameters (see Appendix A.2). With respect to the income levels, we find that $\frac{\partial \tau_i}{\partial y} < 0$ and $\frac{\partial \tau_i}{\partial y^*} > 0$, thus an increase in foreign income, a reduction in the domestic income and hence a widening of the income gap leads to longer migration periods for all migrants. Furthermore, as one would expect intuitively, $\frac{\partial \tau_i}{\partial \gamma_i} > 0$, namely, individuals who have less of a utility discount when consuming abroad display longer migration duration.

Equations (9), (10) and (11) characterise the agents' migration and consumption behavior in the economy and are largely in line with results found

[8] Notice that the consumption patterns established in (10) and (9) also define the savings path, e.g. $s_i^*(t) = y^* - c_i^*$ for $t = 0, ..., \tau_i$.

in the literature following Hill (1987) and Djajic and Milbourne (1988).[9] On this basis, we are now equipped to explore the consequences of agent heterogeneity for aggregated migration patterns.

2.2 Results

Given the above model we can derive results concerning agent heterogeneity, migration decisions and migration flows and stocks.

Permanent and temporary migrants and stayers

The optimal value of τ_i given in (11) may well be larger than an agent's total lifetime, $T_i = 1$. This becomes more likely for very high γ_i, an individual with a small utility discount when living abroad, or for α close to 1 or for a very large income gap $y^* - y$. In fact, what a $\tau_i \geq 1$ implies is that an agent becomes a permanent migrant: the utility value of living and consuming abroad is so large that given the higher income level in the foreign country, returning – even in the last instant of life – creates no additional value. Define by $\bar{\gamma}$ the individual who is indifferent to the question of returning (temporary migration) vs. staying abroad forever (permanent migration). Solving $\tau_i = 1$ from (11) for γ gives the first result:

Lemma 1. *The group of permanent migrants consists of all individuals i with*

$$\gamma_i \geq \bar{\gamma} = \frac{(1-\alpha)y^*}{y^* - \alpha y}\,. \tag{12}$$

[9] One important difference does exist, however: Djajic and Milbourne (1988) find an ambiguous effect of foreign income on the migration duration, since a higher income might result in an earlier return to the home country if the utility function is characterised by a low rate of substitution. This case is excluded here through the specific functional form of utility.

Solving for the consumption volume of a permanent migrant from (9) and (10) after setting $\gamma_i = \bar{\gamma}$, one can verify that $c_i^* = y^*$, i.e. permanent migrants spend their total income in the foreign country and do not save. Furthermore, $\frac{\partial \bar{\gamma}}{\partial y^*} < 0$, such that an increase in the foreign income level – or an increase in the income gap – lowers the threshold value of γ_i, beyond which individuals become permanent migrants.

At the other end of the spectrum, we have those agents who prefer to stay at home instead of migrating. Define by $\underline{\gamma}$ the individual who is completely indifferent to the question of migrating vs. staying at home, i.e. the individual whose optimal migration duration is $\tau_i = 0$. Solving $\tau_i = 0$ from (11) for γ gives:

Lemma 2. *The group of stayers (non-migrants) consists of all individuals i with*

$$\gamma_i \leq \underline{\gamma} = \frac{y - \alpha y^*}{(1-\alpha)y} \,. \tag{13}$$

Since $\gamma_i \in [0,1]$, a necessary condition for at least one individual in the sense of lemma 2 to exist is that $y^* < y/\alpha$. if this condition is violated, then the income gap is so substantial, that all all individuals of the sending country would migrate. For the sake of realism it is assumed that $y^* < y/\alpha$ is fulfilled in the remainder of the Section. Notice that $\frac{\partial \underline{\gamma}}{\partial y^*} < 0$, such that an increase in the foreign income level lowers the threshold value $\underline{\gamma}$, implying that fewer agents are stayers.

It is easy to verify that $\bar{\gamma}$ in (12) is always larger than $\underline{\gamma}$ in (13) as long as the income gap $y^* - y$ is positive. Accordingly, there exists a third group of agents that maximise utility with a $\tau_i \in]0,1[$, i.e. individuals who spend part of their working lives abroad and part at home – temporary migrants. Following the reasoning above, temporary migrants are characterised as follows:

Lemma 3. *The group of temporary migrants consists of all individuals i with*

$$\gamma_i \in]\underline{\gamma}, \bar{\gamma}[\, .$$

Thus, the above results establish that within the population of the home country, three types of agents can be distinguished. While all permanent migrants display identical consumption and migration durations – as do stayers – the group of temporary migrants features varying durations of migration spells. Figure 1 plots the qualities implied by the optimal migration duration from (11) and lemma 1, 2 and 3.

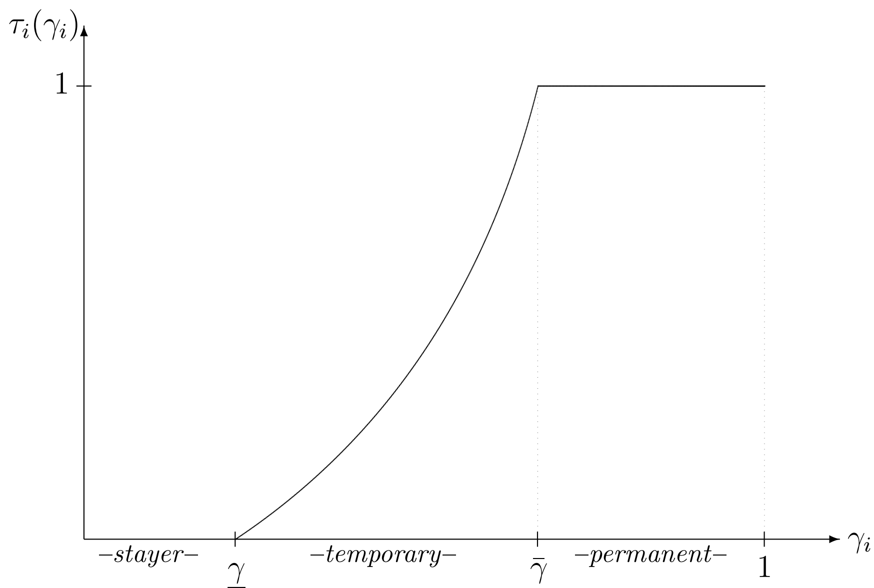

Figure 1: The duration of migration spells for different agents

The aggregate duration of migration

In order to analyse aggregate effects, the distribution of γ_i's in the population must be specified. Here we assume the γ_i's in each cohort to be uniformly distributed on support [0,1]. Accordingly the area under the curve in Figure 1 represents the aggregate duration of all migration spells for a cohort. While τ_i is by definition zero for stayers and one for permanent migrants, Figure 1 shows that the migration duration is monotonically increasing in γ in the interval between $\bar{\gamma}$ and $\underline{\gamma}$.

Integrating (11) over the interval $\underline{\gamma}$ to $\bar{\gamma}$ gives the aggregate duration $\theta^t = \int_{\underline{\gamma}}^{\bar{\gamma}} \tau_i$ of temporary migrants:

$$\theta^t = \alpha N \left(\ln \left(\frac{y^* - y\,\alpha}{y\,(1-\alpha)} \right) - \frac{y^* - y}{y^* - \alpha y} \right). \tag{14}$$

The resulting θ^t captures the total migration duration of all temporary migrants of a cohort, thus in effect both the number of temporary migrants as well as their individual migration durations are captured. As one would intuitively expect, it can be shown that $\frac{\partial \theta^t}{\partial y^*} > 0$, i.e. an increase in foreign income levels has an unambiguously positive effect on the duration of aggregated temporary migration.

Next, under the assumption of uniformly distributed γ_i's, the duration of aggregated permanent migration of a cohort simply becomes

$$\theta^p = 1(1-\bar{\gamma})N = \frac{(y^* - y)\alpha N}{y^* - \alpha y}. \tag{15}$$

Finally, the number of stayers – and since $T_i = 1$ also their aggregate time spent at home – found in a cohort is simply $\underline{\gamma} N = \frac{(y - \alpha y^*)N}{(1-\alpha)y}$.

Combining (14) and (15), we obtain the following result:

Proposition 1. *The aggregate duration of migration, θ, from a single cohort is*

$$\theta = \alpha N \ln \left(\frac{y^* - \alpha y}{y\,(1-\alpha)} \right), \tag{16}$$

- *increases in the foreign income level,* $\frac{\partial\theta}{\partial y^*} > 0$*, and*
- *falls in the domestic income level,* $\frac{\partial\theta}{\partial y} < 0$

Equation (16) arrives at a surprisingly simple specification of the total duration of time spent abroad by the migrants in a given cohort. Lemma 1, 2 and 3 and proposition 1 have clear implications for migration stocks, migration flows and their interaction.

Migration flows and migration stocks

Moving from the migration decisions and durations in a single cohort to migration stocks, one has to specify the number of cohorts coexisting at any instant in time. Let L denote this number. Assuming zero population growth – that is a rate of reproduction of 1 – then the total population at any point in time is LN. Furthermore, assume that each cohort is identical to the previous including their consumption and migration decisions but that descendants' γ_i's are uncorrelated to their parents γ_i's and that reproduction takes place at the end of an agents lifetime.[10] This leads to the following results:

Proposition 2. *The population stocks at every instant in time are*

a) stock of permanent migrants: $S^p = LN(1-\bar{\gamma}) = \frac{LN\alpha(y^*-y)}{y^*-\alpha y}$.

b) stock of temporary migrants: $$S^t = \frac{\theta^t}{N(\bar{\gamma}-\underline{\gamma})} LN(\bar{\gamma}-\underline{\gamma}) = \frac{-\alpha N L}{y^*-y\,\alpha}\left(y^*-y-(y^*-y\,\alpha)\ln\left(\frac{y^*-y\,\alpha}{y\,(1-\alpha)}\right)\right) .$$

[10]This last assumption implies that temporary migrants give birth after they returned to the home country, while only permanent migrants give birth abroad.

c) total stock of migrants: $S = S^p + S^t = \alpha NL \ln\left(\frac{y^* - \alpha y}{y\,(1-\alpha)}\right)$.

and

d) total stock of home population: $H = NL - S(t)$

$$= NL\left(1 - \alpha \ln\left(\frac{y^* - \alpha y}{y\,(1-\alpha)}\right)\right) .$$

Proof (sketch): Proposition 2 a) is the aggregate of all permanent migrants in one cohort times the number of cohorts coexisting at every point in time. Since we have normalized the agent's lifetime to one, this turns out to be $S^p = L\theta^p$. Proposition 2 b) is the average duration of the migration spell of a temporary migrant, $\frac{\theta^t}{N(\bar{\gamma}-\underline{\gamma})}$ times the total number of temporary migrants coexisting at every instant in time, $LN(\bar{\gamma} - \underline{\gamma})$. It follows that $S^t(t) = L\theta^t$. Finally, proposition 2 c) follows from a) and b) and thus $S = L\theta$, and proposition 2 d) follows from c). Thus the stock of migrants is here a fairly simple logarithmic relation of the income gap.

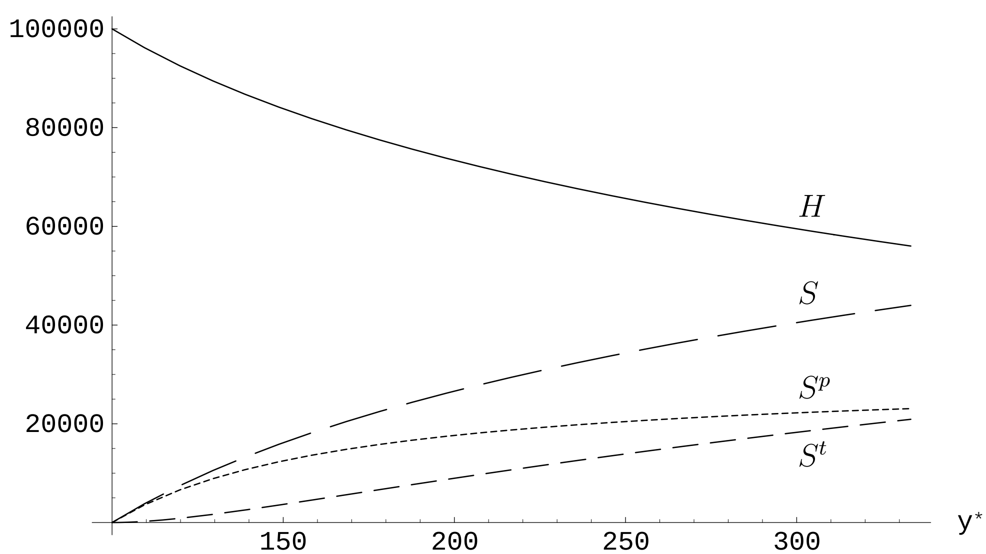

Figure 2: Migration stocks and stock of home population

To see what the relations derived in proposition 2 imply consider Figure 2. Figure 2 plots the stock of temporary, permanent, and total migration as well as the stock of home population as a function of foreign income, y^*. For the actual plot, the following parameter values are employed: $y = 100$, $\alpha = 0.3$, $N = 1000$ and $L = 100$. At $y^* = 100$ there is no income gap and accordingly all agents spend all their working life at home. As the income gap widens, there are initially a few individuals who opt for permanent migration and a few who opt for temporary migration. However, since the income differential is small, the actual amount of time spent abroad is small too. Accordingly from the perspective of the aggregate migration stock, temporary migration contributes relatively little to total migration compared to permanent migration since permanent migrants spend their entire lifetimes abroad. As the income gap widens, the role of temporary migration increases while that of permanent migration decreases.

Finally, consider the migration flows associated with the above stocks, in particular measuring flows occurring during any time interval of length 1. Given that reproduction takes place at the end of an agents life, the number of birth abroad occurring over the time interval 1 are $LN(1-\bar{\gamma})$ $(= S^p)$ while $LN\bar{\gamma}$ birth take place at home, i.e. the entire population has been renewed, however part of any descendant generation are born abroad by migrants.

Proposition 3. *Migration flows over any time interval of length 1 are*

a) gross emigration: $M^e = LN\bar{\gamma}\,(1-\underline{\gamma}) = \frac{LN\alpha y^*(y^*-y)}{y(y^*-\alpha y)}$.

b) gross return migration of home born agents: $M^{r,h} = LN\bar{\gamma}\,(\bar{\gamma}-\underline{\gamma})$

$$= \frac{LN\alpha y^*(y^*-y)^2}{y(y^*-\alpha y)^2} \ .$$

c) gross return migration of foreign born agents: $M^{r,f} = LN(1-\bar{\gamma})\,\bar{\gamma}$

$$= \frac{LN\alpha y^*(y^*-y)(1-\alpha)}{(y^*-\alpha y)^2}.$$

d) net migration: $M = M^e - M^{r,h} - M^{r,f} = 0$

Proof (sketch): Since all individuals that migrate do so at time $t_i = 0$, proposition 3 a) is simply the sum of all home born ($LN\bar{\gamma}$) temporary and permanent migrants. Proposition 3 b) follows from the fact that all cohorts behave identically, such that in equilibrium for every home born temporary outmigrant there is a matching temporary return migrant born in one of the previous cohorts. The return flow of foreign born agents in proposition 3 c) is composed of the share of stayers and temporary migrants ($\bar{\gamma}$) in the total foreign born population ($LN(1-\bar{\gamma})$); or put differently, except for those agents that decide to be permanent migrants all other foreign born individuals will return at some point in time during their life to the home land. Proposition 3 d) follows from a), b) and c).

That the net flow of migrants associated with a given income differential and equilibrium stocks turns out to be zero is driven by the assumption that reproduction takes place at the end of an agents life time. If instead, one assumed that reproduction takes place earlier in an agents life, then some of the temporary migrants would reproduce while staying abroad generating an additional – and unmatched – flow of return migrants that results in *negative* net migration.

Figure 3 plots the migration flows given in proposition 3, for various levels of foreign income, when $y = 100$, $\alpha = 0.3$, $N = 1000$ and $L = 100$.

Corollary 1. *All migration stocks* S^p, S^t *and* S *and the migration flows* M^e, $M^{r,h}$ *and* $M^{r,f}$ *are positive and increasing in the income differential* $y^* - y$.

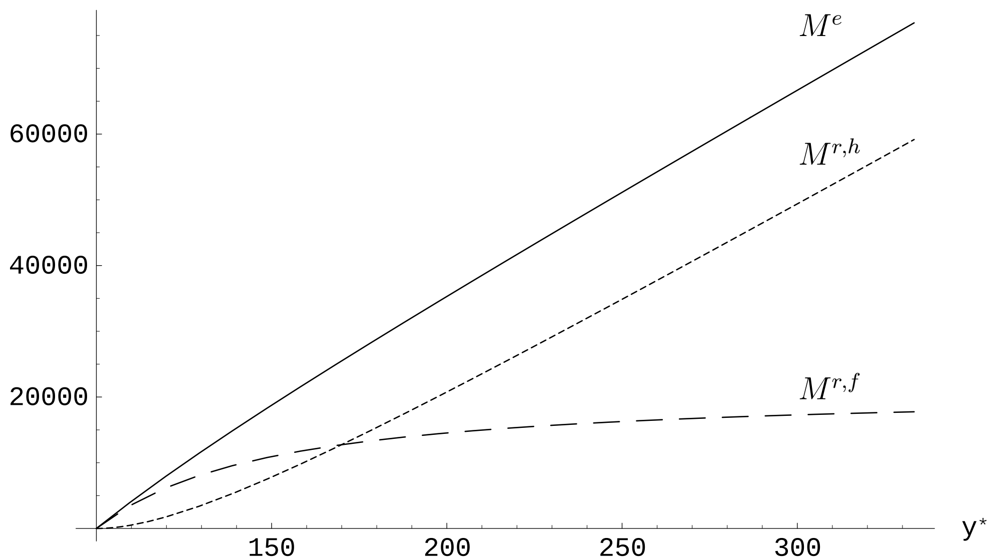

Figure 3: Migration flows in equilibrium

The net migration flow M is zero and independent of the income differential $y^* - y$.

3 Stock vs. flow models

The above results and particular Corollary 1 have important consequences for the empirical estimation of macro migration models. It follows from our model that an equilibrium relationship between the income differential and migration *stocks* but not flows emerges in the long-run. A positive net migration flow can only occur during the transition to some steady state. [11] Yet, most macro migration models in the empirical literature to date state explicitly or implicitly that an equilibrium between migration *flows*

[11] Notice, that this reasoning abstracts from differences in population growth rates among the migrant and the home population and other aspects such as the assimilation and naturalisation of migrants.

and explanatory variables such as the income differential does exists.

We examine these competing hypotheses of the stock and flow models empirically within a cointegration framework. The concept of cointegration is closely related to the notion of equilibrium: a cointegration relationship between variables exists if economic forces drive the system towards the equilibrium defined by the long-run relationship posited (Engle and Granger, 1987). When considering long-run relationships, it becomes necessary to consider the underlying properties of the processes that generate time series variables. If variables follow different stochastic processes over time, spurious regression results can arise that suggest statistically significant long-run relationships between variables, when in fact this is merely evidence of contemporaneous correlations rather than meaningful causal relationships (Granger and Newbold, 1974).

Following the (Engle and Granger, 1987) procedure, we first test for both the stock and the flow model whether the dependent and the explanatory variables are integrated of the same order, and, if this is the case, whether the hypothesis of a cointegration relationship is rejected by our data set. We apply panel unit-root and panel cointegration tests, which increases the statistical power of the tests significantly in comparision to tests based on individual time series. Given a relatively short time dimension in our data set, we use in the final step a dynamic specification for the estimation of the cointegrating vectors and the short-run dynamics.

3.1 Two alternative specifications

The stock model of macro migration implied by Proposition 3, gives for the aggregate migration stock

$$s = \alpha \ln \left(\frac{y^*}{(1-\alpha)y} - \frac{\alpha}{1-\alpha} \right),$$

where s is defined as the share of the migration stock in the total population of the sending country, i.e. $s \equiv S/(LN)$. For empirical purposes, one can approximate the expression for s by

$$s = \beta_0 + \beta_1 \ln \left(\frac{y^*}{y} \right) + \beta_2 \ln(y).$$

We follow furthermore Todaro (1969) and Harris and Todaro (1970) in assuming that income levels are conditioned by employment opportunities in the respective locations. More specifically, if jobs are allocated in each period randomly among the workforce, we can write expected income as the wage times the employment rate, i.e. as $w \times e$. If individuals are risk averse and uncertainty focusses on employment opportunities, it can be expected that the coefficients for the employment variables are larger than those for the wage variables (Hatton, 1995). Moreover, since employment opportunities of migrants in host countries are below those of natives, the coefficient for the employment rate in the host country is larger than that in the source country. Finally, if capital markets are not perfect, liquidity constraints affect migration decisions. Consequently, for a given income difference between the host and the home country, the income level in the source country has a positive impact on migration (Faini and Venturini, 1995).

Based on these considerations we derive the following parsimonious specification for the long-run migration function:

$$s_{it} = a_0 + a_1 \ln\left(\frac{w_{ft}}{w_{it}}\right) + a_2 \ln(w_{it}) + a_3 \ln(e_{ft}) + a_4 \ln(e_{it}) + \nu_{it}, \tag{17}$$

where $i = 1, ..., K$ and $t = 1, ..., T$ are the source country and time indices, s_{it} denotes the migration stock as a percentage of the home population in country i, w_{ft} the wage rate in the host country, w_{it} the wage rate in the home country i, e_{ft} the employment rate in the host country, e_{it} the employment rate in the home country i, and ν_{it} is the error term. The error term is specified as a one-way error component model (Hsiao, 1986), i.e. as $\nu_{it} = \mu_i + \varepsilon_{it}$, where μ_i is a country-specific effect and ε_{it} is white noise.

Compare this to a macro migration model based on migration flows. The standard equation in the empirical literature has the following form (see e.g. Hatton (1995)):

$$\begin{aligned} m_{it} = b_0 + b_1 \ln\left(\frac{w_{ft}}{w_{it}}\right) + b_2 \ln(w_{it}) + b_3 \ln(e_{ft}) + \\ b_4 \ln(e_{it}) + b_5 s_{it} + \nu_{it}, \end{aligned} \tag{18}$$

where m_{it} denotes the net (gross) migration rate as percentage of the home population in country i, i.e. the net flow. The existing stock of migrants is included on the right hand side of the model as a proxy for 'social network' effects which are expected to increase the propensity to migrate by alleviating the adaptation costs in the host country, see Hugo (1981), Massey and Espana (1987), Massey (1990a), Massey (1990b) and Bauer (1995).[12]

The estimation of the migration functions in equations (17) and (18) can

[12] Of course, there exist more possible macro models of migration. The semi-logarithmic functional form has been derived from first principles by (Hatton, 1995), but double-log specifications of macro migration models are common in the literature as well (e.g. Faini and Venturini (1995), Hille and Straubhaar (2001)).

be affected by spurious correlation effects if the regressions involve variables that follow an I(1) or other non-stationary process (see the seminal paper by Granger and Newbold (1974)). The notable exception is the situation when I(1) dependent and explanatory variables form a cointegration set, see Engle and Granger (1987). While there is a general agreement that macroeconomic variables such as income levels and employment rates are rather well represented as I(1) processes, there is still limited evidence on the time series properties of the migration flows and corresponding migrant stock variables. One of the few exceptions in the literature is the Hatton (1995) paper, which provides empirical evidence that all variables in equation (18) are I(1) for UK-US migration from 1870 to 1913, but it is unclear whether this is also supported by other data sets. Particularly puzzling is the fact that the migration flow and the migration stock variable are included in equation (18). Since migration flows can be conceived as (almost) the first difference of migration stocks, they can hardly be I(1) variables if migration stocks are supposed to be I(1) variables as well. Thus, it is reasonable to expect that the migration flow variable is better approximated by an I(0) process if migration stocks are I(1). In this case it is suitable to use the stock model in equation (17) for estimating the long-run migration function.

3.2 Data

A time series analysis of the economic forces which drive international migration requires that migration behaviour is not distorted by institutional or administrative barriers. The EU forms therefore a natural laboratory for students of international migration, since it is the only regional trade area in the world where the free movement of labour and other persons is one of the fundamental freedoms of the common market.[13] The free movement has been fixed already in the Treaty of Rome 1957, and introduced for the six Member States of the then European Economic Community in 1968. In the following decades it has been step by step extended to the 30 members of the EU and European Economic Area (incl. Switzerland), although transitional periods have been applied in the cases of Southern Enlargement of the EU and the present extension of the EU to Central and Eastern Europe.

The sample employed here comprises the migration data from the founding members of the European Community and the three countries from the first Enlargement round (Denmark, Ireland, United Kingdom) to Germany in the period 1973 to 2001. Germany has been chosen as a destination country since it is not only the largest destination of international migration in the EU, but it also reports data on migration stocks and flows since 1967. We begin our analysis in 1973 since this is the year of the first enlargement round. Moreover, the migration data are subject to a visible structural break in 1973 as a consequence of the first oil-price shock. Other events which might have affected migration behaviour such as German unification do not show up in the data as visible structural breaks.

The data on migration stocks and flows come from the German Federal Statistical Office ('Statistisches Bundesamt'). For the stock of migrants, foreign residents as reported by the Central Register of Foreigners ('Ausländerzentralregister') are used as a variable.[14] The stock of foreign residents is reported on December 31 of each year (in some early years on September 30).[15] The number of foreign residents is slightly overstated by

[13]Free labour mobility has been also granted in the Nordic trade area, whose member countries however belong all to EU or the European Economic Area today.

[14]Note that all residents of Germany are obliged to register their place of residence. The figures from the central register of foreigners are based on the reports of the municipalities.

[15]It is sometimes argued that natural population growth and naturalisations distort the

the Central Register of Foreigners, since return migration is not completely registered by the municipalities. Consequently, the figures for the stock of foreign residents has been revised two times following the population censuses of 1972 and 1987. In the econometric analysis, dummy variables are used to control for these statistical breaks.

The data on migration flows stem again from the Central Register of Foreigners. We consider three flow variables: net migration flows, m_{it}, gross inflows, im_{it}, and gross return flows, re_{it}. The migration stock and flow variables are normalised by the population of the home countries, i.e. they are calculated as shares of the corresponding home population. Population figures are depicted from the World Bank's 2002 World Development Indicators and OECD sources. As a proxy for wages and other incomes, the historical series of per capita GDP levels in purchasing power parities from Maddison (1995) has been used. These figures have been extrapolated up to 2001 on basis of the Main Economic Indicators of the OECD. The employment rate is defined as one minus the unemployment rate. Unemployment rates have been taken again from the OECD Main Economic Indicators, and, if not available, complemented by data from national statistical offices. The ILO definition has been used for all unemployment rates.

The descriptive statistics are shown in Table 1.

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migration stock variable. By definition, the increase of the stock of foreign residents equals net immigration plus natural population growth minus the number of naturalisations plus reporting errors for a given period of time. Since our migration variables are calculated as rates, natural population growth cancels out if the rate of natural population growth of migrants equals the rate of natural population growth in the home countries. Thus, if natural population growth of the migrant population in Germany and the source country is similar and the rate of naturalisations is low, the annual increase of the stock of foreign residents as a share in the home population equals almost annual net immigration. Indeed, the annual increase of migration stocks as a share of home population almost equals net immigration rates in our sample.

Table 1 about here

3.3 Testing for unit roots

In the first step of the empirical analysis, the variables are tested for unit roots for making inference on the order of integration. To this end, the Augmented Dickey-Fuller (ADF) test is used for the individual time series and the panel unit root test suggested in Im, Pesaran, and Shin (2003) (IPS-test). The argument for using panel unit root tests instead of univariate unit root tests is that the latter tests are notoriously weak when the root is close to one. In addition, as argued in Shiller and Perron (1985), the problem is aggravated for short time series. Hence, by using the panel data unit root tests, a dramatic increase in terms of power can be achieved (see Levin, Lin, and Chu, 2002).

Tables 2-4 report the results of the ADF and IPS unit-root tests performed on the host- and home-country-specific economic variables. For the IPS-test, the $\bar{t}$-statistic is presented together with the respective critical values, as well as the $w(\bar{t})$-statistic, which is normally distributed (Im, Pesaran, and Shin, 2003). The auxiliary regressions include either an intercept only or an intercept together with a linear deterministic time trend. We present both the results with and without a deterministic trend, since it is not obvious a priori whether the variables considered here exhibit a trending behaviour or not.

Table 2 to 4 about here

As expected, the null hypothesis that the macroeconomic variables, i.e. the relative income ratio and the employment rates, follow I(1) processes, cannot be rejected either in the panel unit root tests or in the majority of the individual ADF tests. Moreover, the null of an I(1) process cannot be rejected for the migrant stock variable either. In contrast, the null of a unit root is clearly rejected for the net and gross migration flow variables in the panel unit root tests. In case of the net and gross migration inflow variable the null of a unit root is rejected by the overwhelming majority of the individual ADF tests, while in case of the gross return migration flow only a minority of the individual ADF tests rejects the null of a unit root. Note that the finding that panel unit root tests clearly reject the null of a unit root for the migration flow variables, while tests for the individual time series do not, is common in the empirical literature (see Wu and Zhang, 1996; Wu, 1996; Papell, 1997).

Thus, the main conclusion from the unit root tests is that the assumption of the standard migration model, that migration *flows* on the one hand, and macroeconomic variables such as GDP per capita levels or employment rates on the other hand, are integrated of the same order, is not supported by the data set employed here. As a consequence, the regression equation is unbalanced as the chosen dependent variable (net or gross migration flows), which has been found to be I(0) variables, is explained by non-stationary I(1) variables.

3.4 Testing for panel cointegration

In order to reconcile the features of the data with the theoretical considerations, the long-run migration function of the migration stock model as specified in equation (17) is employed for the analysis that follows. According to the unit root test results, all the variables of the stock model seem to be I(1), such that they can hypothetically form a cointegration set. Under the assumption of cointegration, the remainder term ϵ_{it} is assumed to be an I(0) variable.

Two specifications of this cointegrating relation are used here: one without a linear deterministic trend and one with. In economic terms, the presence of a linear trend in the regression accounts for the constant growth rate in the migration stock that has been caused by other factors than the income differential and employment conditions. These socioeconomic factors that are not modelled explicitly reflect inter alia different rates of natural population growth in the receiving and the sending countries and decreasing moving costs over time.

Two sets of cointegration tests are reported in Table 5. The first set comprises the results of the two-step Engle-Granger cointegration procedure performed for the variables of every country. The second set comprises the panel cointegration group $t-$test statistics of Pedroni (1999) which aggregates the test statistics obtained in the first place for every country in the panel. For both the specifications without and with trend, the null hypothesis of no cointegration is rejected for 6 out of the 8 countries, albeit in some cases only at the 10% significance level. The more powerful panel cointegration test of Pedroni (1999) rejects the null hypothesis of no cointegration for both model specifications at the 5% significance level.

Table 5 about here

Thus, the results of the cointegration tests suggest that we cannot reject the hypothesis that the variables of the stock model form a cointegrated set.

This allows to estimate the model in equation (17) in order to draw inferences on the parameter values of the cointegrating relations.

4 Estimating the stock model

There are different procedures for estimating both the long-run cointegration relationship and the short-run dynamics. If the variables form a cointegrated set, the cointegrating vector can be consistently estimated in a static regression which completely omits the dynamics of the model (Engle and Granger, 1987). Although the famous super-consistency result (Stock, 1987) indicates that convergence is rather fast, the asymptotic distribution of the least squares estimator and the associated t-statistics is non-normal in finite samples. Moreover, an unadressed 'endogeneity bias' invalidates standard hypothesis testing in samples of finite size.[16] Monte-Carlo evidence suggests that the estimation bias of the cointegrating parameter is smaller in dynamic than in static models (Banerjee, Dolado, Henry, and Smith, 1986). The empirical equation is therefore specified here in form of an error correction model (ECM), which allows estimation of both the long-term cointegrating vector and the short-run dynamics. Note that the ECM has a flexible functional form and imposes few restrictions on the adjustment process.

Specifically, the estimation model has the form

$$\begin{aligned}\Delta s_{it} = \beta_1 s_{i,t-1} + \beta_2 \ln\left(\frac{w_{f,t-1}}{w_{i,t-1}}\right) + \beta_3 \ln(w_{i,t-1}) + \beta_4 \ln(e_{f,t-1}) + \\ \beta_5 \ln(e_{i,t-1}) + \beta_6 \Delta \ln\left(\frac{w_{ft}}{w_{it}}\right) + \beta_7 \Delta \ln(w_{it}) + \beta_8 \Delta \ln(e_{ft}) + \\ \beta_9 \Delta \ln(e_{it}) + \beta_{10} \Delta s_{i,t-1} + \eta' \mathbf{z}_{it} + \mu_i^* + \varepsilon_{it},\end{aligned} \tag{19}$$

[16] See Patterson (2000) for a detailed discussion.

where $\mu_i^* = \mu_i / -\beta_1$ is the long-run value for the country-specific effect, Δ the first-difference operator, $\mathbf{z}_{it}$ a vector of institutional variables and η the corresponding vector of coefficients. Three dummy variables are considered here which should capture the different institutional conditions of migration: guestworker agreements between Germany and the sending country, free movement between the sending country and Germany, and dictatorship in the sending country. The first two variables should cover reduced legal and administrative barriers to migration, the last variable political 'push' factors in the source country. Note that the adjustment parameter of the ECM is given by -β_1, and that the long-term coefficients of the cointegrating relationship are given by -β_k/β_1, where $k = 2, 3...5$. Further lags of the first differences of the dependent variable and lags of first differences of the explanatory variables have not been considered in this specification of the ECM since they appear not significant.

The estimation results of the short-run semi-elasticities of the dynamic model are presented in Table 6. The model is first estimated with a standard fixed effects (within) model (FE). The results of the F-test show that the country-specific effects are indeed highly significant. However, estimating the fixed effects model with ordinary least squares (OLS) may yield inconsistent results if the disturbances are heteroscedastic. One way to obtain a robust covariance matrix is to estimate the model with feasible least squares, which allows for group-wise heteroscedasticity (FGLS(HET)). The likelihood ratio test indeed suggests that the model which allows for group-wise heteroscedasticity is preferable to the homoscedastic model. Finally, spherical disturbances such as common macroeconomic shocks might affect the estimation results. The FGLS(HET+COR) estimator relaxes the assumption of no spherical disturbances by allowing for contemporary correlations across groups. The LR test indicates that the model which allows for both group-wise heteroscedasticity and cross-sectional correlation is preferable to the model which only allows for groupwise heteroscedasticity.[17] Thus, the $FGLS(HET + COR)$ can be expected to yield the most reliable results among the estimators considered here. One caveat is, however, worth noting: the FGLS(HET+COR) estimator tends to understate the standard errors, such that significance levels have to be taken with a grain of salt.

Table 6 about here

Moreover, other objections can be raised against the estimators used here: First, as with all standard panel estimators, the fixed effects estimators are based on the fundamental assumption that the slope parameters are homogeneous. There are good reasons to call this assumption into question. In (Bruecker and Siliverstovs, 2004), the results of various heterogeneous estimators are compared with standard panel estimators in order to shed light on this issue. It can be shown that traditional panel outperform heterogeneous estimators which allow the slope parameters to differ with regard to their forecasting performance (Bruecker and Siliverstovs, 2004). Second, the estimation of dynamic models in samples with a finite time dimension can be affected by simultaneous equation bias, which is caused by the correlation of the lagged dependent variable with the error term. This bias disappears with the time dimension of the panel. It is questionable whether alternative estimation procedures which address the simultaneous equation bias are preferable to standard panel estimators if the time dimension of the panel

[17]All test results are presented in the notes of Table 6.

is substantially larger than the group dimension. Again, it can be shown that the forecasting errors of GMM estimators which address this problem is larger than that of traditional panel estimators (Bruecker and Siliverstovs, 2004).

Before interpreting the estimation results, recall that according to the theoretical considerations discussed in Section 3.1, positive signs are expected for the difference in per capita GDP levels, the per capita GDP level in the sending country, and the employment rate in the host country, and negative signs for home employment rates since they increase employment opportunities in the source country.

The estimation results confirm these expectations: first, the coefficient for the lagged migration rate is negative and highly significant. Note that this further supports the stock model, since most flow models expect a positive coefficient for (lagged) migration stocks. However, the coefficients for the lagged first difference of the migration stock is positive and again highly significant in all three regressions. This can be interpreted as evidence of so-called network or 'herd effects' (Epstein and Hillman, 1998). In the long run, however, the propensity to migrate in the sending countries decreases as the share of the population already living abroad increases.

Second, both for the income differential and home income, we find the expected positive coefficients. In most regressions, these effects are highly significant. One exception is the FGLS(HET) model, where the income differential is only significant at the 10% level.

Third, the employment rate in the receiving country is highly significant in all regressions and its coefficient is substantially larger than the coefficient for the income difference. This highlights the importance of labour market conditions in the host countries and confirms the expectations outlined in

Section 3.1. The employment rate in the sending countries has the expected negative coefficient and is significant in all three regressions. However, this parameter has a much smaller value than the employment rate in the receiving country, which again supports the expectations stated in Section 3.1. Note that many empirical studies find that home employment opportunities have been insignificant or have actually increased migration (see Greenwood (1975) for a review). The results here do not confirm these findings, but they do show that home employment has a much weaker impact than employment in the receiving countries. One possible explanation for this phenomenon is that favorable employment opportunities in home countries might have ambiguous effects on migration, since higher employment rates reduce incentives to seek employment abroad on the one hand, while on the other they help to lift liquidity constraints that potential migrants face, and in doing so, may encourage migration.

Fourth, the variables in first differences again have the expected signs, but do not appear significant in all regressions. Given the rather short time dimension of the panel, this is not surprising.

Fifth, the institutional variables have the expected signs in all regressions, but we observe substantial differences in the size of the parameters and their significance. The coefficient of the guestworker dummy is large and appears highly significant in all three regressions. In contrast, the size of the parameter for the free movement dummy is only one-tenth of that of the guestworker dummy. Moreover, it only appears significant in the FGLS(HET+COR) regression, there however at the 1% level. The rather low impact of the free movement dummy might be explained by the rather low variance in the sample: in past accession rounds, free movement was only granted either to countries with a similar or higher per capita income to the existing EU

Members (Austria, Denmark, Finland, Iceland, Norway, Sweden, UK), or to countries where the stock of migrants was already very large and presumably close to equilibrium levels (Greece, Ireland, Portugal, Spain). Thus, the variance in the sample might be too low to detect the impact of free movement from past Enlargement episodes. Finally, the estimated parameter for the dictatorship dummy appears to be large and highly significant. This result highlights the well-known fact that political push-factors have an important impact on migration and can easily dominate economic forces.

To sum up, these results do all indicate the model of migration stocks to be both of empirical relevance and to yielding sound results.

5 Conclusion

In this paper, we examined the macro determinants of migration both from a theoretical and an empirical perspective. The model presented in the theoretical part establishes a long-run equilibrium, in which individuals can stay their entire life in the home country, migrate temporarily abroad – with individual durations of migration spells – or stay permanently in a foreign country depending on their preferences. This model generated insights in the mechanics of migration stocks and flows. The number of migrants, the duration of migration spells and therewith the stock of migrants all increases with the income difference between the host and the home country, while net migration ceases to zero. The gross emigration and return migration rates are related to the stocks of permanent and temporary migrants. Since the stock of permanent and temporary migrants and are a positive function of the income differential. Consequently, existing empirical migration models, estimating net migration flows, instead of stocks, may be misspecified.

In the empirical part of the paper the determinants of international migration have been analysed within a cointegration framework. The methodological aspects of the analysis can be summarised as follows: first, the results of the panel unit-root and panel cointegration test suggest that the standard flow migration model is misspecified – at least for the data set used here. The traditional migration model in the empirical model explains migration flows by a number of explanatory variables such as GDP per capita, (un-)employment rates, (lagged) migration stocks and institutional variables. It is widely acknowledged in the literature that macroeconomic variables such as GDP and employment are non-stationary variables, or, more specifically, I(1) variables. The existence of a long-run equilibrium between migration flows and the traditional set of macroeconomic variables requires therefore that migration flows are I(1) as well. The tests carried out in the empirical part of this paper, however, indicate that migration rates are stationary, while migration stocks are I(1) variables. Moreover, the empirical analysis carried out here suggests that the hypothesis of a cointegration relationship between migration stocks and the explanatory variables cannot be rejected for our data set. This can be interpreted as empirical support for the theoretical hypothesis that migration stocks and explanatory variables such as the income differential and employment variables form an equilibrium relationship.

Our findings have some important policy consequences. The flow model suggests that migration does not stop before expected income levels between host and source countries have converged to a certain threshold level, which is determined by the costs of migration. In case of persistent differences in expected income levels, either the total population will eventually migrate or migration will not occur in the first place. In contrast, the stock model predicts that migration ceases when the benefits of migration equal the costs to the marginal migrant, such that a long-run equilibrium between migration stocks and expected income emerges. Consequently, migration may cease despite the existence of large income differences.

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A Appendix

A.1 First-order conditions of the Lagrangian of the migrant's maximisation problem

Define by L the Lagrangian for the maximisation problem in (3) under the budget constraint (4). The first-order conditions are:

$$\frac{\partial L}{\partial c_i^*(t)} = \alpha\gamma_i^{1-\alpha}{c^*}_i^{-(1-\alpha)} - \lambda = 0\,, \tag{A.1}$$

$$\frac{\partial L}{\partial c_i(t)} = \alpha c_i^{-(1-\alpha)} - \lambda = 0\,, \tag{A.2}$$

$$\frac{\partial L}{\partial \tau_i} = \gamma_i^{1-\alpha} c_i^*(t)^\alpha + c_i(t)^\alpha - \lambda(c_i^*(t) - c_i(t) + y - y^*) = 0\,, \tag{A.3}$$

$$\frac{\partial L}{\partial \lambda} = \tau_i y^* + (1-\tau_i)y - \tau_i c_i^*(t) - (1-\tau_i)c_i(t) = 0\,, \tag{A.4}$$

where λ is the shadow value of wealth.

A.2 Derivatives of the optimal τ_i

Differentiating (11) with respect to y^*, y and γ_i gives:

$$\frac{\partial \tau_i}{\partial y^*} = \frac{y - y\,\alpha}{(y^* - y)^2} > 0\,, \tag{A.5}$$

$$\frac{\partial \tau_i}{\partial y} = \frac{y^*\,(\alpha - 1)}{(y^* - y)^2} < 0\,, \tag{A.6}$$

$$\frac{\partial \tau_i}{\partial \gamma_i} = \frac{\alpha}{(1-\gamma)^2} > 0\,. \tag{A.7}$$

Table 1: Descriptive statistics

variable	obs.	mean	standard deviation	minimum	maximum
s_{it}	594	0.8510	0.9860	0.0370	4.5650
m_{it}	594	0.0002	0.0008	-0.0053	0.0073
im_{it}	594	0.0009	0.0012	0.0001	0.0107
$\ln(w_{ft}/w_{it})$	594	0.2420	0.3740	-0.3820	1.5880
$\ln(w_{it})$	594	9.4550	0.4250	7.7760	10.3470
$\ln(e_{ft})$	33	-0.0580	0.0290	-0.0990	-0.0060
$\ln(e_{it})$	594	-0.0660	0.0490	-0.2770	0.0000
$GUEST_{it}$	594	0.0340	0.1810	0	1
$FREE_{it}$	594	0.5640	0.4960	0	1
$DIKT_{it}$	594	0.0490	0.2160	0	1

Table 2: Unit-root test results (s_{it}, m_{it})

	$\mathbf{s}_{it}$				$\mathbf{m}_{it}$			
	without		*with trend*		*without*		*with trend*	
	test-stat.	lags	test-stat.	lags	test-stat.	lags	test-stat.	lags
AUS	-3.082**	0	-2.742	0	-3.648***	1	-4.519***	1
BEL	-1.474	0	-2.846	1	-3.441**	1	-4.109**	1
DK	-0.375	1	-2.631	1	-3.754**	3	-4.116**	1
ESP	-1.223	1	-2.018	1	-4.745***	1	-4.664***	1
FIN	-0.770	1	-2.145	1	-2.606	0	-2.660	0
FRA	-0.895	0	-2.434	0	-3.101**	1	-3.752**	1
GRE	-2.894*	4	-2.842	4	-3.849***	1	-3.814**	1
ICE	-0.957	0	-1.510	0	-2.722*	0	-2.737	0
IRE	-0.742	0	-2.422	1	-2.674*	0	-2.651	0
ITA	-2.463	1	-2.390	1	-4.073***	0	-4.652***	1
LX	-2.457	0	-3.368*	0	-2.292	1	-2.205	1
NET	-2.125	0	-3.307*	0	-3.321**	1	-3.913**	1
NOR	-0.733	0	-1.750	0	-2.476	0	-2.446	0
POR	-2.436	1	-2.478	1	-2.804*	0	-2.823	0
SWE	0.236	0	-1.291	0	-3.686***	1	-3.649**	1
SWI	-1.482	0	-2.855	0	-4.033***	1	-5.141***	1
TK	-2.736*	1	-1.587	1	-3.156**	1	-3.598**	1
UK	-2.065	0	-1.358	1	-3.017**	1	-3.669**	1
IPS-Test	-0.36076		-0.84187		-8.3439***		-7.216***	

**, *, * denote the rejection of the H_0 of a unit root at the 1%, 5%, and 10% significance level, respectively.– Critical values of the ADF-test for the the rejection of the H_0-hypothesis of a unit-root are -3.70, -2.98 and -2.62 at the 1%, 5% and 10% significance level, respectively, in the regressions without trend, and -4.32, -3.57 and -3.22 at the 1%, 5%, and 10% significance level, respectively, in the regressions with deterministic trend (at 32 observations).– Critical values of the IPS test for rejection of the H_0 of a unit root are -2.32, -1.64, -1.28 at the 1%, 5%, and 10% significance level, respectively.

Table 3: Unit-root test results (im_{it}, $\ln(w_{ft}/w_{it})$)

	$\mathbf{im}_{it}$				$\mathbf{ln(w_{ft}/w_{it})}$			
	without trend		*with trend*		*without trend*		*with trend*	
	test-statistic	lags	test-statistic	lags	test-statistic	lags	test-statistic	lags
AUS	-1.875	2	-5.249***	1	-2.984**	0	-3.278*	0
BEL	-3.413**	1	-3.477*	1	-2.057	0	0.080	5
DK	-2.091	1	-2.429	1	-2.975**	1	-3.153	1
ESP	-2.440	2	-1.954	2	-1.600	1	-0.861	1
FIN	-2.342	1	-2.131	1	-2.903*	1	-3.362*	1
FRA	-3.383**	1	-3.714**	1	-2.040	1	-1.305	0
GRE	-1.793	2	-1.939	2	-2.361	0	-2.818	0
ICE	-1.530	0	-1.549	0	-3.182**	1	-2.594	0
IRE	-1.594	0	-1.254	0	-4.720	0	1.282	0
ITA	-3.359**	2	-3.790**	0	-1.513	4	-1.513	4
LX	-1.697	0	-1.412	0	1.266	0	-1.202	0
NET	-2.035	2	-1.505	2	-1.414	0	-1.114	0
NOR	-2.250	1	-2.202	1	-1.989	1	-3.171	1
POR	-1.916	0	-1.961	0	-1.647	0	-3.334*	1
SWE	-1.815	1	-2.257	0	-2.007	0	-2.697	1
SWI	-2.478	2	-1.307	2	-1.167	0	-3.430*	3
TK	-1.626	0	-3.485**	1	-2.239	0	-2.531	0
UK	-3.270**	1	-3.107	1	-3.338**	1	-2.123	1
IPS-Test	-3.611**		-1.806**		-0.619		0.406	

**, *, * denote the rejection of the H_0 of a unit root at the 1%, 5%, and 10% significance level, respectively.– Critical values of the ADF-test for the the rejection of the H_0-hypothesis of a unit-root are -3.70, -2.98 and -2.62 at the 1%, 5%, and 10% significance level, respectively, in the regressions without trend, and -4.32, -3.57 and -3.22 at the 1%, 5%, and 10% significance level, respectively, in the regressions with deterministic trend (at 32 observations).– Critical values of the IPS-Test for rejection of the H_0 of a unit-root are -2.32, -1.64, -1.28 at the 1%, 5%, and 10% significance level, respectively.

Table 4: Unit-root test results (ln (w_{it}), ln (e_{it}), ln(e_{ft}))

	ln (w_{it})				**ln (e_{it})**		**ln(e_{ft})**	
	without trend		*with trend*		*without trend*		*without trend*	
	test-statistic	lags	test-statistic	lags	test-statistic	lags	test-statistic	lags
AUS	-3.307**	0	-2.738	3	-1.812	2	-	
BEL	-1.819	0	-3.376*	4	-1.888	1	-	
DK	-0.872	0	-2.380	0	-1.921	1	-	
ESP	-0.611	1	-3.379*	1	-1.627	2	-	
FIN	-0.819	1	-2.993	1	-1.566	2	-	
FRA	-1.831	0	-2.946	2	-1.664	1	-	
GRE	-2.742*	0	-1.882	3	-1.338	1	-	
ICE	-3.538**	0	-2.856	1	-2.051	1	-	
IRE	4.653	0	1.402	0	-1.319	1	-	
ITA	-1.399	0	-1.742	1	-1.339	1	-	
LX	1.187	0	-1.223	0	-0.982	3	-	
NET	-1.041	0	-2.597	1	-1.788	1	-	
NOR	-2.131	1	-2.554	1	-1.814	1	-	
POR	0.572	4	-5.090***	3	-1.795	4	-	
SWE	-1.061	0	-2.465	1	-2.240	1	-	
SWI	-1.318	0	-2.984	1	-1.051	1	-	
TK	-1.704	0	-2.409	0	-2.808	1	-	
UK	-0.755	0	-3.536*	1	-1.607	2	-	
GER	-		-		-		-1.306	2
IPS-Test	2.307		-1.129		-0.945		-	

**, *, * denote the rejection of the H_0 of a unit root at the 1%, 5%, and 10% significance level, respectively.– Critical values of the ADF-test for the the rejection of the H_0-hypothesis of a unit-root are -3.70, -2.98 and -2.62 at the 1%, 5%, and 10% significance level, respectively, in the regressions without trend, and -4.32, -3.57 and -3.22 at the 1%-, 5%- and 10%-significance level, respectively, in the regressions with deterministic trend (at 32 observations).– Critical values of the IPS-Test for rejection of the H_0 of a unit-root are -2.32, -1.64, -1.28 at the 1%, 5%, and 10% significance level, respectively.

Table 5: Cointegration test results

	without trend		*with trend*	
	test-statistic	lags	test-statistic	lags
AUS	-5.359***	0	-4.993***	0
BEL	-3.084	4	-3.453	1
DK	-2.653	1	-3.319	0
ESP	-5.629***	3	-5.500***	3
FIN	-3.853	1	-3.645	1
FRA	-2.801	1	-3.083	0
GRE	-3.380	4	-4.4906***	3
ICE	-3.452	3	-3.964*	3
IRE	-3.211	2	-2.476	1
ITA	-2.640	2	-3.824	3
LX	-4.302***	4	-4.112*	4
NET	-3.543	0	-3.522	0
NOR	-2.750	0	-3.378	3
POR	-3.039	4	-3.163	1
SWE	-2.430	0	-2.640	0
SWI	-3.721	0	-5.003***	1
TK	-2.737	1	-2.996	2
UK	-3.282	2	-3.529	2
Group *t*-statistics	-1.588*		-1.515*	

**, *, * denote the rejection of the H_0 of a unit root in the residuals at the 1%, 5%, and 10% significance level, respectively. In the regressions with intercept, the critical values for the rejection of the H_0 of a unit root are -4.73, -4.11 and -3.83 at the 1%, 5%, and 10% significance level, respectively.– In the regressions with intercept and deterministic trend, the critical values are -4.65, -4.16 and -3.84 at the 1%, 5%, and 10% significance level respectively. See Hamilton (1994).– The group t-statistic has the asymptotic standard normal distribution. The one-sided critical values for the rejection of the H_0 of a unit root in the residuals are -2.63, -1.64 and -1.28 at the 1%, 5%, and 10% significance level, respectively.

Table 6: Estimation results

	FE[1]			**FGLS (HET)**[2]			**FGLS (HET+COR)**[3]		
	coeff.		t-stat.	coeff.		t-stat.	coeff.		t-stat.
$s_{i,t-1}$	-0.150	***	-6.06	-0.126	***	-8.66	-0.143	***	-20.78
$\ln(w_f/w_i)_{t-1}$	0.087	**	2.21	0.042	*	1.92	0.084	***	5.32
$\ln(w_i)_{t-1}$	0.104	***	3.57	0.056	***	3.40	0.099	***	6.43
$\ln(e_f)_{t-1}$	0.733	***	3.36	0.342	***	3.65	0.613	***	7.03
$\ln(e_i)_{t-1}$	-0.163	*	-1.95	-0.106	**	-2.31	-0.131	***	-11.77
$\Delta\ \ln(w_f/w_i)_t$	0.102	**	2.33	0.037		0.33	0.120		1.11
$\Delta\ \ln(w_i)_t$	0.358	***	2.99	0.184		1.57	0.282	**	2.61
$\Delta\ \ln(e_f)_t$	0.851	***	3.22	0.408	*	1.83	0.548	**	2.62
$\Delta\ \ln(e_i)_t$	-0.225		-0.97	-0.164		-1.37	-0.163	***	-5.70
$\Delta\ s_{i,t-1}$	0.411	***	4.75	0.302	***	7.89	0.410	***	19.03
$GUEST_{it}$	0.098	***	6.27	0.105	***	5.69	0.109	***	11.64
$FREE_{it}$	0.008		0.91	0.000		0.07	0.006	***	3.69
$DIKT_{it}$	0.062	**	2.01	0.012		0.77	0.048	***	5.91
$STAT(1972)$	-0.112	***	-2.97	-0.048	***	-7.56	-0.101	***	-15.86
$STAT(1987)$	-0.083	***	-3.77	-0.048	***	-7.79	-0.082	***	-13.59
adjusted R^2	0.61			-			-		
Log-Likelihood	-			1280			1661		

1) Fixed Effects (within) regression. The $F(17{,}543)$ statistic for the H_0 that all $\mu_i = 0$is 9.80^{***}.– 2) Feasible Generalised Least Squared (FGLS) regression with country dummies. The robust estimation of the covariance matrix allows for groupwise heteroscedasticity in the disturbances. The $\chi^2(17)$ statistic for the LR test of the heteroscedastic vs. the homoscedastic model is 761.04^{***}.– 3) *FGLS* regression with country dummies. The robust estimation of the covariance matrix allows for groupwise heteroscedasticity in the disturbances and correlation across groups. The $\chi^2(33)$ statistic for the LR test of the heteroscedastic and correlated vs. the heteroscedastic model is 762.58^{***}.