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The European Crisis and Migration to Germany:

Expectations and the Diversion of Migration Flows

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Abstract

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Keywords: international migration; multiple destinations; diversion; dynamic discrete choice model; expectations.

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The European Crisis and Migration to Germany: Expectations and the Diversion of Migration Flows

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Abstract

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1   Introduction

What has been the impact of the economic crisis that began in 2008 on migration in Europe? The answer to such a question may seem evident, as the asymmetric effects of the crisis upon different countries have been matched by a transformation of the landscape of migration flows within Europe, with Spain and Germany representing two polar cases in this respect. Spain, which had experienced a surge in the share of immigrants in its population from 4 percent in the late 1990s to 14 percent in 2007 (Bertoli and Fernández-Huertas Moraga, 2013), recorded *negative* net migration flows in 2011 and 2012 (INE, 2013), while Germany experienced a substantial increase in net migration flows, which stood at 0.5 percent of its population in 2012 compared to 0.1 percent in the decade before (Statistisches Bundesamt, 2013), with only a minor part of this surge that can be directly traced back to the European countries which have been more severely affected by the crisis.[1] This, in turn, suggests that providing a convincing answer to our initial question requires to account for the *diversion* of migration flows determined by the crisis, as worsening labor market conditions and dismal economic prospects in some European countries are likely to have induced prospective migrants to adjust their destination choices. Consider, for instance, Romania, which is the origin country that experienced the largest increase in gross migration flows to Germany in recent years: the size of this increase corresponds roughly to the size of the decline in migration flows from Romania to Spain, as revealed by Figure 1.

The standard econometric approach in the international migration literature is ill-equipped to answer our research question, as it is based on an underlying modeling of migration as a forward-looking but permanent decision, which is not adjusted or reversed as the attractiveness of alternative destinations changes over time, as observed by Kennan and Walker (2011). The micro-foundation of the decision to migrate rests on the canonical random utility maximization model (McFadden, 1974, 1978), with the (either implicit or explicit) assumption that the deterministic component of utility corresponds to "the present value of expected earnings at destination" (Ortega and Peri, 2013, p. 55). This, in turn, leads to estimate bilateral migration rates as a function of the differential in *current* attractiveness of the destination and of the origin country only, thus greatly constraining the scope for the

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[1]Germany recorded a net immigration flow of 70,000 persons from Greece, Italy, Portugal and Spain in 2012, while the corresponding flow from the new EU member states in Central and Eastern Europe stood at 190,000 persons (Statistisches Bundesamt, 2013).

existence of diversion effects. If prospective migrants consider the possibility of temporary migration[2] and of making additional moves in the future, then the decision to stay or migrate today also reflects such an option (Burda, 1995), so that the expectations about the future attractiveness of alternative destinations could also be influencing current location choices. Coming back to our example, the expectation of a prolonged period of high unemployment in Spain might contribute to explain the recent increase in migration flows from Romania to Germany.

We draw on recent contributions to the economic literature (Artuç *et al.*, 2010; Kennan and Walker, 2011; Arcidiacono and Miller, 2011) to propose a dynamic discrete choice model that accounts for the sequential nature of the decisions to migrate, where the deterministic component of utility for each destination in period t depends on the expected value of the optimal sequence of location choices from period $t + 1$ onwards. This model, which is consistent with the forward-looking and path-dependent nature of migration decisions, can be solved analytically under standard distributional assumptions on the stochastic component of utility, but the corresponding choice probabilities differ starkly from those generated by a canonical RUM model. Specifically, the assumption that the stochastic component of utility follows and identically and independently distributed Extreme Value Type-1 distribution (McFadden, 1974) does *not* allow us to write the logarithm of the odds ratio of two locations solely as a function of the current attractiveness of the two locations, as the odds ratio depends also upon the future attractiveness of all alternatives in the choice set, and on the whole structure of bilateral migration costs. We label this dependency as *dynamic* multilateral resistance to migration, to differentiate it from the one that arises when introducing more general distributional assumptions on the canonical RUM model (Bertoli and Fernández-Huertas Moraga, 2013), which we refer to as *static* multilateral resistance to migration.[3]

We derive the expression for the bias that arises when bilateral migration rates are estimated only as a function of the characteristics of the origin and of the destination country, following a long-established tradition in the migration literature (Hanson, 2010). In terms of our model, this traditional estimation approach can be justified only if we assume either that

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[2]See, *inter alia*, Djajić and Milbourne (1988), Dustmann and Kirchkamp (2002), Dustmann (2003) and Brücker and Schröder (2012) for models of temporary migration.

[3]Arcidiacono and Miller (2011) show how unobserved individual heterogeneity can be incorporated into a dynamic RUM model.

individuals behave myopically, abstracting from the future consequences of current location choices, or that there are no migration costs.

We demonstrate that the Common Correlated Effects, CCE, estimator proposed by Pesaran (2006) allows to control for dynamic multilateral resistance to migration when estimating the determinants of bilateral migration flows with aggregate data. Kennan and Walker (2011) use their dynamic RUM model to estimate the determinants of internal migration decisions with individual-level longitudinal data, so that a contribution of our paper is to demonstrate that a sequential model of migration can be estimated with a less data-demanding approach.[4] We adopt the same econometric approach followed by Bertoli and Fernández-Huertas Moraga (2013), so that this is able to remove the bias due to multilateral resistance to migration, irrespective of its dynamic or static origin. In our framework, we consider explicitly expectations on the time-varying attractiveness of alternative destinations, thus departing from the estimation approaches adopted by Artuç *et al.* (2010) and Arcidiacono and Miller (2011).

Our sequential model of migration is used to analyze the determinants of migration flows from the member states of the European Economic Association, EEA,[5] to Germany based on a high-frequency administrative dataset from January 2006 to December 2012.[6,7] This paper is, to the best of our knowledge, the first one to explicitly consider how expectations on the future attractiveness of alternative destinations can lead to a diversion of current migration flows. The EEA represents an area with unique institutional features, as its legislation favors the free mobility of workers between its member states, thus facilitating repeated moves by the migrants. Sequential moves are not consistent with a representation of the location-decision problem that potential migrants face through a canonical discrete choice model,

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[4]Artuç *et al.* (2010) also use aggregate data to identify the switching costs that workers in the US face when changing the sector they are employed in, but their estimation approach does not deal with general forms of individual unobserved heterogeneity, what we can call static multilateral resistance to migration.

[5]The EEA encompasses the whole EU plus Iceland, Liechtenstein and Norway; although Switzerland is not *de jure* a member of the EEA, it has ratified a series of bilateral agreements with the EU that allows to usually regard it as a *de facto* EEA member state.

[6]The migration data have a monthly frequency; other papers using monthly or quarterly migration data in an econometric analysis are Hanson and Spilimbergo (1999), Orrenius and Zavodny (2003) and Bertoli and Fernández-Huertas Moraga (2013).

[7]The CCE estimator has satisfactory small sample properties already for the longitudinal and cross-sectional dimension of our data according to the Monte Carlo simulations in Pesaran (2006).

and this strengthens the case for the adoption of the dynamic model of migration decisions that underpins our estimation approach.

We also provide evidence on the direct role played by expectations on the decision to migrate by augmenting the vector of determinants of location-specific utility with a forward-looking variable, which can reflect the expectations about future economic prospects at origin held by potential migrants. More specifically, we use the yields on the secondary market of government bonds with a residual maturity of 10 years as a proxy for future economic conditions. This choice is supported by the evidence that we provide using data from 14 waves of the Eurobarometer survey that concerns about personal job market prospects and economic conditions in general in the year to come are closely related to the evolution of the 10-year bond yields.[8]

Our econometric analysis reveals that variations in the unemployment rate at origin significantly influence the bilateral migration rate to Germany, but that the size of this effect is grossly overestimated in standard specifications that do not control for dynamic multilateral resistance to migration. This bias, whose direction is consistent with the one implied by our sequential model of migration, is removed once we resort to the CCE estimator, which reveals that the elasticity of the bilateral migration rate with respect to unemployment at origin stands at 0.5. We also provide evidence that a 10 percent increase in the 10-year bond yields at origin is associated with a 1.4 percent increase in the bilateral migration to Germany, significantly below the (biased) estimate that we get when we do not account for multilateral resistance to migration. The standard estimation approaches that do not fully account for the forward-looking and path-dependent nature of the decision to migrate can produce biased estimates of the determinants of international migration flows. Our estimates reveal that the bias in the estimated effect of the unemployment rate at origin is not solely due to its correlation with the domestic future unemployment rate, but also to its correlation with the future attractiveness of alternative destinations.

This paper is related to four main strands of literature. First, the literature on the determinants of international migration flows (Clark *et al.*, 2007; Pedersen *et al.*, 2008; Lewer and den Berg, 2008; Mayda, 2010; Grogger and Hanson, 2011; Beine *et al.*, 2011; Belot and

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[8]A key feature of this variable is that it certainly belongs to the information set upon which potential migrants take their decisions, as the media coverage of the yields of 10-year bonds has substantially increased in recent years when the crisis unfolded; Farré and Fasani (2013) demonstrate that information on fundamental economic variables in the media significantly impacts migration decisions.

Hatton, 2012; Bertoli *et al.*, 2011; Belot and Ederveen, 2012; McKenzie *et al.*, 2013; Beine *et al.*, 2013), and more specifically to the papers that have relaxed the distributional assumptions on the underlying RUM model (Ortega and Peri, 2013; Bertoli *et al.*, 2013; Bertoli and Fernández-Huertas Moraga, 2012, 2013), and those that have analyzed the determinants of migration to Germany, mainly in the context of the EU's Eastern enlargement (Vogler and Rotte, 2000; Boeri and Brücker, 2001; Fertig, 2001; Flaig, 2001; Sinn *et al.*, 2001; Brücker and Siliverstovs, 2006). Second, the literature on discrete choice models (McFadden, 1974, 1978; Small and Rosen, 1981; Cardell, 1997; Wen and Koppelman, 2001; Train, 2003; de Palma and Kilani, 2007). Third, this also paper draws on the papers that have proposed dynamic discrete choice models (Pessino, 1991; Artuç *et al.*, 2010; Kennan and Walker, 2011; Arcidiacono and Miller, 2011; Bishop, 2012; Artuç, 2013). Fourth, the literature on the estimation of linear models with a common factor structure in the error term (Pesaran, 2006; Bai, 2009; Pesaran and Tosetti, 2011).

The remainder of the paper is structured as follows: Section 2 presents a RUM model that describes the sequential location-decision problem that potential migrants face, and it derives the equation to be estimated. Section 3 introduces our sample and data sources, and it provides empirical evidence that supports our reliance on 10-year bond yields as proxies for the expectations about future economic conditions at origin. Section 4 contains the relevant descriptive statistics, and Section 5 presents the results of our econometric analysis. Section 6 draws the main conclusions of the paper.

2 A sequential model of migration

We consider a set of agents, each of them denoted by i, located in country j that have to choose their preferred location from a set of countries D, which includes n elements, for each period $t = 1, ..., T$. The expected utility of opting for country k at time t is given by:

$$U_{ijkt} \equiv w_{kt} - c_{jk} + \beta V_{t+1}(k) + \epsilon_{ikt} \tag{1}$$

This depends on (i) a deterministic instantaneous component w_{kt}, (ii) a deterministic time-invariant component c_{jk} that describes the cost of moving from j to k,[9] (iii) the

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[9]We assume that bilateral migration costs are time-invariant, but the model can be readily extended to allow for an exogenous evolution over time in migration costs.

discounted value, with time discount factor $\beta \leq 1$, of the expected utility $V_{t+1}(k)$ from optimally choosing the preferred location from time $t+1$ onwards conditional upon being in k at time t, and on (iv) a stochastic individual- and time-specific component ϵ_{ikt}. For simplicity, we assume that there is no uncertainty about the evolution over time of the deterministic component of utility w_{jt} for all $j \in D$,[10] and that potential migrants also know all possible bilateral migration costs c_{jk}. We also assume that individual i chooses her preferred location after having observed the realizations of the stochastic component of utility at time t for all countries, but without any information on their future realizations. This, in turn, explains why we refer to U_{ijkt} in (1) as expected utility, as $V_{t+1}(k)$ is a random variable. Notice that $w_{kt} + \beta V_{t+1}(k)$ does *not* represent the present value of expected instantaneous utility in country k, as the continuation payoff $V_{t+1}(k)$ also reflects the value of the option to move away from k at some time $s \geq t+1$. Hence, our model differs from a RUM model where the deterministic component of utility is interpreted as the present value of expected earnings (or instantaneous utility more generally).[11]

2.1 The continuation payoff $V_{t+1}(k)$

The continuation payoff $V_{t+1}(k)$ depends on k as individuals located in different countries can face a different vector of bilateral migration costs. In the absence of migration costs, then $V_{t+1}(k) = V_{t+1}$ for any location k chosen at time t.[12]

We can obtain an analytic expression for the value of the continuation payoff by backward induction, and introducing distributional assumptions on the stochastic component of utility. Specifically, let us focus first on $V_T(k)$. We have that the continuation payoff for the last period is given by the inner product of a vector $\boldsymbol{p}_{kT}$ that describes the probability of moving from k to any location in D at time $t = T$ and of a vector $\boldsymbol{u}_T$ that describes the expected utility from choosing each location in D at time $t = T$, conditional upon the fact that this

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[10]This assumption is actually unnecessary, as discussed by Artuç *et al.* (2010).

[11]Notice that in such a model not only the decision to migrate is permanent, but also the decision to stay cannot be reversed, as observed by Kennan and Walker (2011).

[12]$V_{t+1}(k)$ would still depend on k even in the absence of migration costs if we allowed the time-varying deterministic component of utility w_{kt} in (1) to vary across individuals and to depend on their past migration history; this more general version of the model could, for instance, handle a positive return to the time elapsed since migration, introducing a greater persistence in location choices over time, but would not alter the key insights from our sequential migration model.

is the utility-maximizing alternative. Formally:

$$V_T(k) \equiv \boldsymbol{p}_{kT}{}' \boldsymbol{u}_T$$

If we assume that the stochastic component of utility follows an independent and identically distributed EVT-1 distribution (McFadden, 1974), then we have that:

$$\boldsymbol{p}_{kT} = \left(\sum_{l \in D} e^{w_{lT} - c_{kl}} \right)^{-1} \begin{pmatrix} e^{w_{1T} - c_{k1}} \\ \dots \\ e^{w_{nT} - c_{kn}} \end{pmatrix}$$

Furthermore, de Palma and Kilani (2007) demonstrate that these distributional assumptions imply that the expected utility does *not* vary across alternatives, and Small and Rosen (1981) already provided the analytical expression for the expected value from the choice situation, so that:

$$\boldsymbol{u}_T = \left[\gamma + \ln \left(\sum_{l \in D} e^{w_{lT} - c_{kl}} \right) \right] \boldsymbol{1}$$

where γ is the Euler's constant and $\boldsymbol{1}$ is a $n \times 1$ vector whose elements are all equal to 1. This allows us to rewrite $V_T(k)$ as follows:[13]

$$V_T(k) = \gamma + \ln \left(\sum_{l \in D} e^{w_{lT} - c_{kl}} \right)$$

Then, the expected utility from locating in country k at time $t = T - 1$ can be rewritten as:

$$U_{ijkT-1} = w_{kT-1} - c_{jk} + \beta \ln \left(\sum_{l} e^{w_{lT} - c_{kl}} \right) + \beta \gamma + \epsilon_{ikt}$$

If we go one more step back, to define the expected continuation payoff for period $t = T - 1$, we can observe that:[14]

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[13]Notice that $V_T(k)$ is an increasing and convex function of w_{lT}, for any $l \in D$.

[14]We rely here on $w_{kT-1} - c_{lk} + \beta V_T(k)$ as the (expected) deterministic component of the attractiveness of country k at time $t = T - 1$, which includes also the discounted value of the continuation payoff from time $t = T$, as in Kennan and Walker (2011).

$$V_{T-1}(l) = \gamma + \ln \left(\sum_{k \in D} e^{w_{kT-1} - c_{lk} + \beta V_T(k)} \right)$$

More generally, we can rewrite the expression for location-specific utility at time t as follows:

$$U_{ijkt} = w_{kt} - c_{jk} + \beta \left[\gamma + \ln \left(\sum_{l \in D} e^{w_{lt+1} - c_{kl} + \beta V_{t+2}(l)} \right) \right] + \epsilon_{ikt} \qquad (2)$$

2.2 Choice probabilities at time t

The distributional assumptions on the stochastic component entail that the vector of choice probabilities $\boldsymbol{p}_{kt}$, for any $t = 1, ..., T$, can be written as:

$$\boldsymbol{p}_{kT} = \left(\sum_{l \in D} e^{w_{lT} - c_{kl} + \beta V_{t+1}(l)} \right)^{-1} \begin{pmatrix} e^{w_{1T} - c_{k1} + \beta V_{t+1}(1)} \\ \dots \\ e^{w_{nT} - c_{kn} + \beta V_{t+1}(n)} \end{pmatrix}$$

If we take the logarithm of the ratio of the probability to opt for country k over the probability to stay in country j in period t, then we get:[15]

$$\ln \left(\frac{p_{jkt}}{p_{jjt}} \right) = w_{kt} - c_{jk} - w_{jt} + \beta \left[V_{t+1}(k) - V_{t+1}(j) \right] \qquad (3)$$

The expression in (3) depends on (i) the difference between the deterministic component of utility at time t in k and in j only, and (ii) on the difference in the discounted value of the expected continuation payoff from k and j.[16]

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[15]We normalize the cost of staying in j to zero, i.e., c_{jj}=0; as choice probabilities depend only on the difference in utility across countries rather than on their levels, the normalization is immaterial.

[16]Artuç *et al.* (2010) and Arcidiacono and Miller (2011) go one step further, and express the difference between the continuation payoffs $V_{t+1}(k) - V_{t+1}(j)$ as a function of future choice probabilities, exploiting the fundamental result of Hotz and Miller (1993); we do not follow this approach, as we are also interested in providing direct evidence on the role of expectations in shaping current location decisions; our estimates on the effect of the unemployment rate at origin are robust to the adoption of their proposed estimation approach.

2.3 The standard specification in the literature

A long-standing tradition in the international migration literature is to express the logarithm of the ratio of the probability to opt for country k over the probability to stay in country j in period t as depending only on the attractiveness of the two countries (Hanson, 2010). What are the assumptions that would allow us to rewrite (3) in such a way?

We would need to assume either that (i) individuals take myopic decisions, i.e., $\beta = 0$, or that (ii) there are no migration costs, i.e., $c_{jk} = 0$ for any $j, k \in D$. Assumption (i) deprives the continuation payoffs $V_{t+1}(k)$ and $V_{t+1}(j)$ in (3) of any relevance for current location decisions, while assumption (ii) entails that $V_{t+1}(k) = V_{t+1}(j)$, so that the solution to the sequential location decision problem is memoryless, as it is independent from past choices.

Needless to say, either of the two assumptions is highly implausible: assuming that individuals are myopic is starkly at odds with the representation of migration as a forward-looking investment decision (Sjaastad, 1962), while the absence of migration costs stands in sharp contrast with the empirical evidence that the scale of international migration flows is severely constrained by policy-induced migration costs (Pritchett, 2006; Clemens, 2011; Mayda, 2010; Ortega and Peri, 2013; Bertoli and Fernández-Huertas Moraga, 2012). Hence, it is important to gain a better understanding of the implications of the dependency of current location decisions on $V_{t+1}(k)$ and $V_{t+1}(j)$.

2.4 The future attractiveness of alternative destinations

Let us rewrite the expression for the continuation payoff $V_{t+1}(k)$:

$$V_{t+1}(k) = \gamma + \ln \left(\sum_{l \in D} e^{w_{lt+1} - c_{kl} + \beta V_{t+2}(l)} \right)$$

If we derive it with respect to the attractiveness of a country h at time $s = t + 1$, we get:

$$\frac{\partial V_{t+1}(k)}{\partial w_{ht+1}} = p_{kht+1} \tag{4}$$

A marginal variation in w_{ht+1} induces a change in the continuation payoff $V_{t+1}(k)$ that is equal to the probability of moving from k to h at time $t + 1$. We can also observe that:

$$\frac{\partial^2 V_{t+1}(k)}{\partial w_{ht+1} \partial c_{kh}} = -p_{kht+1}(1 - p_{kht+1})$$

The impact on $V_{t+1}(k)$ of a variation in the attractiveness of destination h at time $t+1$ is larger the lower are the bilateral migration costs from k to h. Similarly, for $s = t + 2$, we have that:[17]

$$\frac{\partial V_{t+1}(k)}{\partial w_{ht+1}} = \sum_{l \in D} \frac{\partial V_{t+1}(k)}{\partial V_{t+2}(l)} \frac{\partial V_{t+2}(l)}{\partial w_{ht+2}} = \beta \sum_{l \in D} p_{klt+1} p_{lht+2} = \beta \pi_{kh}(t, t+2) \qquad (5)$$

where $\pi_{kh}(t, t+2)$ represents the probability of moving from country k at time t to country h at time $t+2$. We can easily generalize (4) and (5) to any time $s = t + 1, ...T$ as follows:

$$\frac{\partial V_{t+1}(k)}{\partial w_{hs}} = \beta^{s-t-1} \pi_{kh}(t, s)$$

This eventually allows us to write the impact on $V_{t+1}(k)$ of a *permanent* variation in the future attractiveness of country h as follows:

$$\sum_{s=t+1}^{T} \frac{\partial V_{t+1}(k)}{\partial w_{hs}} = \sum_{s=t+1}^{T} \beta^{s-t-1} \pi_{kh}(t, s) > 0 \qquad (6)$$

The expression in (6) allows us to compute the impact of a permanent variation of the attractiveness of country h on the logarithm of the ratio of the choice probabilities at time t. Specifically, we have that:

$$\sum_{s=t+1}^{T} \frac{\partial \ln(p_{jkt}/p_{jjt})}{\partial w_{hs}} = \sum_{s=t+1}^{T} \beta^{s-t} [\pi_{kh}(t, s) - \pi_{jh}(t, s)] \gtrless 0 \qquad (7)$$

The term within the summation is given by the difference in the probabilities of moving respectively from k and j at time t to h at time s. This, in turn, depends on the probability of all paths of length $s - t$ that optimally lead an individual i to move from either k or j to country h. Hence, the sensitivity of current location decisions with respect to a variation in the future attractiveness of an alternative destination does *not* depend only on the bilateral migration costs c_{kh} and c_{jh}, but on the whole structure of bilateral migration costs, as migrants can make an indirect move from, say, k to h via (at most) $s - t - 1$ countries. This also entails that we cannot, in general, sign (7) without imposing a structure on the matrix of bilateral migration costs.

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[17]Notice that the variation in $V_{t+1}(k)$ depends on the variation in the expected value of the choice situation for $t+2$ for all countries in the choice set.

The partial derivative in (7) allows us to conclude that, unless we are willing to assume that $\beta = 0$ or there are no migration costs, our sequential model of migration is characterized by multilateral resistance to migration (Bertoli and Fernández-Huertas Moraga, 2013), as the logarithm of the ratio of choice probabilities at time t is sensitive to variations in the future attractiveness of alternative destinations. This occurs even though we have assumed, as most of the literature does, that the stochastic component of location-specific utility is i.i.d. EVT-1: this assumption suffices to make (3) independent from the current attractiveness of alternative destinations, but this logarithm of the ratio of the choice probabilities remains dependent on the future attractiveness of alternative destinations. In our model, multilateral resistance to migration does *not* arise because of more general distributional assumptions as in Bertoli and Fernández-Huertas Moraga (2013), but rather because we accounted for the sequential nature of the location-decision problem that individuals face. This is why we refer to it as *dynamic* multilateral resistance to migration.

2.5 Estimation

Assume that we have an empirical counterpart for the logarithm of the ratio of the choice probabilities, and let it be denoted by y_{jkt}. Assume also, as the literature does, that the deterministic component of the attractiveness of location-specific utility can be expressed as a linear function of a vector of variables $\boldsymbol{x}$, so that:

$$y_{jkt} = \boldsymbol{\alpha}' \left(\boldsymbol{x}_{jkt} - \boldsymbol{x}_{jjt} \right) + r_{jkt} + \eta_{jkt} \tag{8}$$

where $\boldsymbol{x}_{jkt}$ and $\boldsymbol{x}_{jjt}$ reflect respectively the attractiveness at time t of country k and j for an individual located in j at time $t-1$, $r_{jkt} = \beta \left[V_{t+1}(k) - V_{t+1}(j) \right]$ is the term that captures the influence on y_{jkt} on the future attractiveness of all countries, and η_{jkt} is a well-behaved error term.

The term r_{jkt} is a non-linear function of the (time-varying) future attractiveness of all countries belonging to the choice set, and it also depends on the vector of parameters α to be estimated. We can rely on (7) to provide a linear approximation of r_{jkt}, which can be expressed as:

$$r_{jkt} \approx \widetilde{r}_{jk} + \boldsymbol{\gamma}_{jk}' \boldsymbol{f}_t \tag{9}$$

where:

$$\boldsymbol{\gamma}_{jk} = \begin{pmatrix} \sum_{s=t+1}^{T} \beta^{s-t} \left[\pi_{k1}(t,s) - \pi_{j1}(t,s) \right] \\ \ldots \\ \sum_{s=t+1}^{T} \beta^{s-t} \left[\pi_{kn}(t,s) - \pi_{jn}(t,s) \right] \end{pmatrix} \bigg| \boldsymbol{x}_{jks} = \widetilde{\boldsymbol{x}}_{jk}, \, \forall k \in D$$

and:

$$\boldsymbol{f}_t = \begin{pmatrix} \boldsymbol{x}_{j1t} - \widetilde{\boldsymbol{x}}_{j1} \\ \ldots \\ \boldsymbol{x}_{jnt} - \widetilde{\boldsymbol{x}}_{jn} \end{pmatrix}$$

with $\widetilde{\boldsymbol{x}}_{jk}$ representing the average of the determinants attractiveness of k for individuals coming from j over the period of analysis, and $\widetilde{r}_{jk}$ is the dynamic multilateral resistance to migration term evaluated in correspondence to these average values for all countries. This approximation of r_{jkt} allows us to rewrite (8) as follows:

$$y_{jkt} = \boldsymbol{\alpha}' \left(\boldsymbol{x}_{jkt} - \boldsymbol{x}_{jjt} \right) + \widetilde{r}_{jk} + \boldsymbol{\gamma}_{jk}' \boldsymbol{f}_t + \eta_{jkt} \tag{10}$$

and it suggests relying, as in Bertoli and Fernández-Huertas Moraga (2013), on the Common Correlated Effect, CCE, estimator proposed by Pesaran (2006) to deal with the threat to identification posed by multilateral resistance to migration. Specifically, Pesaran (2006) demonstrates that a consistent estimate of $\boldsymbol{\alpha}$ can be obtained when the common factors $\boldsymbol{f}_t$ are serially correlated and correlated with the vectors $\boldsymbol{x}_{jkt}$ and $\boldsymbol{x}_{jjt}$ from the estimation of the following regression:

$$y_{jkt} = \boldsymbol{\alpha}' \left(\boldsymbol{x}_{jkt} - \boldsymbol{x}_{jjt} \right) + \alpha_{jk} d_{jk} + \boldsymbol{\lambda}_{jk}' \widetilde{\boldsymbol{z}}_t + \eta_{jkt} \tag{11}$$

where d_{jk} are dyadic fixed effects and the vector of auxiliary regressors $\widetilde{\boldsymbol{z}}_t$ is formed by the cross-sectional averages of the dependent and of all the independent variables. The consistency of the estimates is established by Pesaran (2006) by demonstrating that $\boldsymbol{\lambda}_{jk}' \widetilde{\boldsymbol{z}}_t$ converges in quadratic mean to $\boldsymbol{\gamma}_{jk}' \boldsymbol{f}_t$ as the cross-sectional dimension of the panel goes to infinity, with the longitudinal dimension being either fixed or also diverging to infinity.[18] Section 5 provides further details on the exact specification of the equation that will be estimated.

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[18]See Eberhardt *et al.* (2013) for a non-technical introduction to the CCE estimator.

2.5.1 The standard estimation approach in the literature

What happens if we rely on an estimation approach that does *not* control for the multilateral resistance to migration term in (8)? Such a standard approach is going to give rise to a biased and inconsistent estimate of $\boldsymbol{\alpha}$, which cannot be interpreted as reflecting the structural parameters of the underlying RUM model. Specifically, this occurs whenever the current determinants of the attractiveness of country k and j, $\boldsymbol{x}_{jkt}$ and $\boldsymbol{x}_{jjt}$, are correlated with the future attractiveness of the two countries, or of any alternative destination in the choice set. When the confounding influence of r_{jkt} is not controlled for, so that the multilateral resistance to migration term ends up in the error term, we have that this correlation determines the endogeneity of all the elements in the vector of regressors.[19]

Imagine, for the sake of concreteness, that the attractiveness of a country depends on its current unemployment rate, and that an increase in the rate of unemployment in country j is positively correlated with its future level in country j itself, and in some alternative destinations. In such a case, the estimated coefficient of the unemployment rate is biased if the confounding influence of multilateral resistance to migration is not controlled for, as it also reflects the influence of variations in the future attractiveness of some countries on current location decisions. Clearly, this represents a relevant threat to identification in our case, as we will be focusing on a set of European countries that also represented relevant destinations for other countries in the region and that have been experiencing an economic crisis with relevant shared component over the past few years. In such a case, the direct effect of, say, a rise in unemployment in Italy on migration flows to Germany can be confounded by the simultaneous surge of the Spanish unemployment rate, which might have diverted the flow of Italian migrants from Spain to Germany. Differently from the source of *static* multilateral resistance to migration described by Bertoli and Fernández-Huertas Moraga (2013), information about the prevailing patterns of correlation in the data does *not* suffice to sign the direction of the ensuing bias, unless we are willing to introduce assumptions on the structure of bilateral migration costs; this is why we write that a persistent worsening in labor market conditions *might* have diverted the Italian migration flows towards Germany.[20]

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[19]The endogeneity due to multilateral resistance to migration implies that the approach proposed by Driscoll and Kraay (1998) to deal with the non-spherical error term in (8) cannot be applied here, as it rests on the assumption of the exogeneity of the regressors.

[20]This ambiguity will disappear in our estimates, where the structure of fixed effects enables us to sign the expected direction of the bias due to dynamic multilateral resistance to migration irrespective of the

2.5.2 A focus on expectations

Let us consider the derivative of the logarithm of the ratio of the choice probabilities with respect to the future attractiveness of the country of origin j:

$$\sum_{s=t+1}^{T} \frac{\partial \ln\left(p_{jkt}/p_{jjt}\right)}{\partial w_{js}} = \sum_{s=t+1}^{T} \beta^{s-t}\left[\pi_{kj}(t,s) - \pi_{jj}(t,s)\right] < 0 \tag{12}$$

As long as migration costs are positive,[21] we always have that the probability of staying in j at time t is higher than the probability of moving to j from any other country, i.e., $\pi_{jj}(t,s) > \pi_{kj}(t,s)$. This, in turn, allows to conclude that an improvement in the future attractiveness of country j unambiguously reduces the logarithm of the ratio of the current probability to migrate from j to k over the corresponding probability of staying in j.

If, at time t, we have data about a vector of variables $\boldsymbol{q}_{jt}$ which is informative about the attractiveness of the origin country j for $s \geq t+1$, then we could augment the equation to be estimated (11) with $\boldsymbol{q}_{jt}$:[22]

$$y_{jkt} = \boldsymbol{\alpha}'\left(\boldsymbol{x}_{jkt} - \boldsymbol{x}_{jjt}\right) + \alpha_{jk}d_{jk} + \boldsymbol{\phi}'\boldsymbol{q}_{jt} + \boldsymbol{\lambda}_{jk}'\widetilde{\boldsymbol{z}}_t + \eta_{jkt} \tag{13}$$

While the estimation of (11) allows us to control for the confounding effect of the future attractiveness of the countries in the choice set on current location choices, the estimation of (13) would also allow us to directly estimate the influence of the future attractiveness of the origin country on the current size of bilateral migration flows.

2.5.3 Extensions

The proposed sequential model of migration can be extended in a number of possible directions, which do not alter the main insights for the estimation that we have derived from it.

First, we could introduce an additional source of uncertainty in the model, beyond the one represented by the individual-specific stochastic component of utility in (1). Specifically, we could consider macroeconomic uncertainty, by relaxing the assumption that individuals

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structure of bilateral migration costs.

[21]Recall that we have normalized c_{jj} to zero, for any $j \in D$.

[22]Similarly, we can augment with $\boldsymbol{q}_{jt}$ the standard specification, which does not control for multilateral resistance to migration.

know the evolution of the attractiveness of each country in the choice set. If we treat w_{ks}, with $s = t + 1, ..., T$ and $k \in D$ as a random variable, then the expected value of the continuation payoff $V_s(k)$ would be computed as an integral over the whole distribution of the future attractiveness of the n countries.

Second, we could follow Arcidiacono and Miller (2011) and Bishop (2012), generalizing the distributional assumptions on the stochastic component of utility in (1) along the lines of Bertoli and Fernández-Huertas Moraga (2013).[23] This would, in turn, imply that the logarithm of the ratio of the choice probabilities in (3) would also become sensitive to the current, and not just to the future, attractiveness of alternative destinations, thus combining in a single model both what we called static and dynamic multilateral resistance to migration.[24]

3 Sample composition and data sources

This section describes the sample of origin countries included in our analysis, together with the data sources for the migration data and for the other variables.

3.1 Sample

The sample of origin countries included in our analysis is composed by all member states of the European Economic Association, EEA, plus Switzerland. The EEA includes all member states of the European Union, EU, together with Iceland, Liechtenstein and Norway, and it represents an area of free mobility of labor.[25] It also extends to Switzerland, which has not joined the EEA but has signed *de facto* equivalent bilateral agreements with the EU.[26] The only exceptions are represented by Liechtenstein and Malta, as the migration data that we use do not provide figures on migration flows from Liechtenstein to Germany, and the

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[23]For instance, the literature on discrete choice models provides us with an analytical expression for the continuation payoff $V_{t+1}(k)$ when the distributional assumptions on ϵ_{ikt} give rise to a nested logit model (Train, 2003).

[24]Both types of multilateral resistance to migration call for the same estimation approach to handle them, as discussed in Bertoli and Fernández-Huertas Moraga (2013) and in Section 2.5 above.

[25]See Part III of the Agreement of the European Economic Area, Official Journal No. L 1, January 3, 1994, and later amendments.

[26]See http://eeas.europa.eu/switzerland/index_en.htm (last accessed on December 12, 2012); we will at times slightly abuse the legal definitions, referring to the EEA as if it also includes Switzerland.

series for Malta contains some zero entries.[27] This sample includes 28 countries of origin, slightly below the threshold of 30 for which Pesaran (2006) provides Monte Carlo evidence on the correct size of the CCE estimator. This is why, as a robustness, we also consider an extended sample including two major non-EEA countries of origin, namely Turkey and Croatia, whose citizens do not benefit from the same rules concerning free mobility.[28]

3.2 Data sources

3.2.1 Migration data

The data on gross migration inflows are provided by the Federal Statistical Office of Germany (Statistisches Bundesamt, 2013).[29] The Federal Statistical Office reports monthly data series on arrivals of foreigners by country of origin since January 2006.[30] We use all the observations that are currently available, namely from January 2006 until December 2012, which gives us 84 monthly observations for each one of the countries in our sample.

The German migration figures are based on the population registers kept at the municipal level. Registration is mandatory in Germany, as stated by the German registration law approved in March 2002 ("Melderechtsrahmengesetz").[31] This law prescribes that each individual has to inform the municipality about any change of residence. The law does *not* subordinate the need to register to a minimum duration or to the scope of the stay, though there are exceptions for foreign citizens whose intended duration of stay in Germany is below two months, so that tourists do not have to register.[32] Figures are reported separately for German and foreign citizens. Foreigners are defined as all individuals who do not possess the

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[27]As we weight observations by population at origin in our estimates, the exclusion of these two countries from the sample is immaterial, as they jointly represent less than 0.1 percent of the population of the EEA.

[28]Turkish immigrants represent the largest migrant community in Germany, but total inflows have been rather moderate in recent years; Croatia is, together with Serbia, the main migrant-sending country among former Yugoslavian countries, but recent inflows have been also relatively modest.

[29]This is the same data source as in OECD (2012).

[30]The country of origin is defined as the country where an individual was resident before moving to Germany.

[31]The data are collected at the end of each month and reported about six weeks later by the municipalities to the local statistical offices of the Federal States and to the Federal Statistical Office; see Statistisches Bundesamt (2010) for an in-depth outline of this dataset.

[32]Further exceptions are allowed for diplomats or foreign soldiers and their relatives who do not have to register.

German citizenship according to Article 116(1) of the German constitutional law ("Grundge-setz"), which also encompasses stateless persons. The inflows of the so-called ethnic Germans ("Spätaussiedler") are reported together with the inflows of German citizens.

This administrative data source provides us with an accurate information on bilateral migration flows to Germany, as migrants have an incentive to register, and municipalities also have an incentive to accurately update their population registers.[33] Specifically, registration is a necessary precondition to obtain the income tax card that is required to sign any employment contract,[34] including for seasonal work, and to issue an invoice if self-employed.[35] Also, landlords usually require a proof that their would-be tenants have registered. Furthermore, the municipalities have an incentive to record new residents properly since their tax revenues depend on the number of registered inhabitants, so that fees are levied against the persons who do not comply with the mandatory registration.[36]

This data source gives us $28 \times 84 = 2,352$ observations for our main sample, with inflows representing 62.5 percent of total gross inflows of migrants to Germany over our seven-year period of analysis.

3.2.2 Other variables

We draw the information on the mid-year size of the population at origin, which is used for defining our dependent variable[37] and to weight the observations in our sample, from

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[33]Notice that the same immigrant might be recorded in the data more than once in case of repeated migration episodes.

[34]The limited incidence of informal employment in Germany suggests that the number of illegal migrants not covered by this administrative data source is likely to be small, and all the more so for the origin countries included in our sample since the main irregular communities are estimated to come from Turkey, Afghanistan and Iraq. No EEA country is among the top ten of the largest irregular communities according to these estimates (Schneider, 2012; Vogel and Assner, 2011).

[35]The self-employed also need to register in order to set up an address for their firm.

[36]The administrative data also contain information on monthly outflows of foreigners, based on cancelation from the local population registers; these are less reliable than inflow data, as migrants have fewer incentives to de-register upon departure.

[37]As it is common in the literature, this the logarithm of the ratio between the gross flow of migrants from j to k at time t over the size of the total population at origin at time t; this definition drives a wedge with the theoretical model, as the denominator of the ratio should actually be represented by the portion of the population that chose to stay at origin at time t, while the total population also includes immigrants and returnees. We provide evidence below that this proxy of the theoretically relevant concept does not

World Bank (2013) and from EUROSTAT (2013c). The latter data source also provides information on the size of the population by five-year age cohorts, that we will use to perform some robustness checks on our estimates.

The location-specific expected utility corresponding to the country of origin is explicitly modeled as a function of (various lags) of the unemployment rate, GDP per capita and the yields on 10-year government bonds.[38] Furthermore, the econometric analysis allows the bilateral migration rate to Germany to depend also on relevant immigration policy variables and on a number of dyadic factors that are controlled for but whose effects are not identified (see Section 5).

The data for the monthly rate of unemployment for all countries in the sample but Switzerland come from EUROSTAT (2013b), while the Swiss unemployment rate were obtained from Statistik Schweiz (2010). The series, which are based on the ILO definition of unemployment, are seasonally adjusted. The data for real quarterly GDP are derived from the International Financial Statistics of the IMF (2013); when the original series are not seasonally adjusted, we adjust them following the method proposed by Baum (2006). We rely on population figures from World Bank (2013) to obtain real GDP per capita series.

The third key variable in our analysis is represented by the yields on the secondary market of government bonds with a residual maturity of 10 years. For EU countries, the primary data source is represented by the European Central Bank, with the ECB series being available at EUROSTAT (2013a) and the OECD (2013). We complemented these data sources with data from National Central Banks. The ECB does not provide 10-year bond yields figures for Estonia, as the country has a very low public debt financed with bonds of a shorter maturity.[39] To fill this gap in the data, we have regressed the 10-year bond yields on a linear transformation of the sovereign ratings from Fitch (2013), and used the estimated coefficients from this auxiliary regression to predict the 10-year bond-yields for Estonia.[40] As a robustness check, we also exclude Estonia from the sample, to ensure

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influence our estimates. See also Section 5 for a discussion on this point.

[38]All the independent variables have been collected since January 2005, as we will be using an optimally selected number of lags for the independent variables.

[39]The ECB states that "there are no Estonian sovereign debt securities that comply with the definition of long-term interest rates for convergence purposes. No suitable proxy indicator has been identified." (source: http://www.ecb.int/stats/money/long/html/index.en.html, last accessed on December 12, 2012).

[40]The estimation of the relationship between 10-year bond yields and sovereign ratings includes country fixed effects; still, the inclusion of origin dummies in our analysis of the determinants of migration flows

that the imputation of the 10-year bond yields does not affect our estimates.[41]

Finally, we defined two dummy variables for the accession of Bulgaria and Romania to the EU in January 2007, and for the concession of free movement of labor to Germany in May 2011 to the citizens of eight countries that accessed the EU in 2004,[42] and that had been subject to transitional agreements that partly limited their right to work in other member states.

3.3   Ten-year bond yields and expectations

The yields that prevail on the secondary market for government bonds with a residual maturity of 10 years represent a usual focal point along the curve that relates yields to maturity, which is commonly reported in the media and plays a key role in European treaties.[43] Differentials in bond-yields within the EEA, and in particular within the Eurozone, are mainly caused by fiscal vulnerabilities,[44] and by the perceptions about the risk of default, the liquidity in the sovereign bonds markets and the time-varying risk preferences of investors (Barrios *et al.*, 2009).[45] Movements in the spreads can have significant consequences, as a rise in sovereign yields tend to be accompanied by a widespread increase in long-term interest rates faced by the private sector (the so-called sovereign ceiling effect), affecting both investment and consumption decisions. On the fiscal side, higher government bond yields

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entails that we do not use between-country variability for identification, so that the level of the predicted Estonian bond yields is actually irrelevant.

[41]The same procedure has been used to predict 10-year bond yields for the two countries in our extended sample, as bond-yields were missing for Turkey in 2005 and for Croatia over the whole period.

[42]Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovak Republic and Slovenia; Cyprus and Malta also joined the EU in 2004, but the right to the free movement of labor was granted to the citizens from these two countries without any transitional periods.

[43]The Article 121 of the Treaty establishing the European Community states that "the durability of convergence achieved by the Member State and of its participation in the exchange-rate mechanism of the European Monetary System being reflected in the long-term interest-rate levels" (Official Journal of the European Communities C 325/33, December 24, 2002), and the European Central Bank gathers harmonized data on 10-year bonds to assess convergence on the basis of Article 121.

[44]In general, the yields of the 10-year government bonds reflect (*i*) the expectations about future interest rates, (*ii*) inflation and (*iii*) the risk premium required by the investors; in what follows, we implicitly assume that point (*iii*) is driving the evolution, across time and space, of the 10-year bond yields in the EEA.

[45]For the broad literature which analyses the economic determinants of the spread in interest rates see, *inter alia*, von Hagen *et al.* (2011), Bernoth and Erdogan (2010) and Caggiano and Greco (2012).

imply higher debt-servicing obligations when the debt is rolled over (Caceres *et al.*, 2010), which can, in turn, induce the implementation of austerity programs to stabilize debt ratios that can further depress economic conditions (Blanchard and Leigh, 2013).

This is why we can presume that the evolution of the 10-year bond yields can be correlated with the evolution of the expectations held by the citizens about the future economic outlook of their own country, which can, in turn, influence their decisions to migrate.

3.3.1 The Eurobarometer survey

The hypothesis that 10-year government bond yields capture individual expectations on personal economic prospects is proved here based on the Eurobarometer surveys. The Eurobarometer surveys are based on approximately 1,000 interviews conducted in European countries twice a year since 1973.[46] We selected the waves and the countries corresponding to the sample of countries that we use in our main econometric analysis.

We thus drew the data from all the 14 waves of the Eurobarometer survey conducted between the Spring 2006 and the Fall 2012 in 27 countries.[47] We focused on the question: "what are your expectations for the year to come: will [next year] be better, worse or the same, when it comes to your personal job situation?", and we analyzed the determinants of the share of respondents who expect their job situation to worsen over the next year. Notice that the data from the survey cannot be directly used in the estimation of the determinants of bilateral migration rates, as they have a lower frequency than the other variables and they do not cover all the countries in our sample.

Table 1 presents some descriptive statistics for the 14 waves of the Eurobarometer survey for the 27 countries listed above.[48] There is notable variability across countries in the share of respondents that expect their personal job situation to worsen over the next year, varying from an average of 2.9 percent for Denmark to 26.7 percent for Hungary. Interestingly,

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[46]The exceptions with respect to the sample size are represented by Germany (1,500 individuals), Luxembourg (600) and United Kingdom (1,300).

[47]The countries are Austria, Belgium, Bulgaria, Cyprus, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Latvia, Lithuania, Luxembourg, Netherlands, Poland, Portugal, Romania, Slovak Republic, Slovenia, Spain, Sweden and the United Kingdom; Iceland is included only since 2010, while two countries in our sample (Norway and Switzerland) are not covered by the Eurobarometer survey.

[48]German data are not used in the analysis, which is restricted to the 26 origin countries that belong to our sample.

Germany is the country that experienced the largest decline in this share between the first (April 2006) and the last wave (November 2012) of the survey, down from 12 to 5 percent. All other countries but five experienced an opposite pattern, with the share of respondents who expressed their concern that increased by as much as 27 percentage points in Greece.[49] As expected, although there are differences across countries, Table 1 delivers the image of the European dimension of the crisis. Importantly, the correlation in economic conditions across the countries in our sample creates a threat to the econometric analysis of the determinants of bilateral migration flows to Germany that is discussed in Section 2 and addressed by our identification strategy.

Needless to say, we do not claim that a simple multivariate analysis can unveil a causal relationship between the expectations about the future labor market conditions and the interest rate on the sovereign rate bond, as the latter may well respond to concerns about the economic perspectives of a country. What we are interested in is to uncover whether the interest rate on 10-year government bonds is positively associated with expectations on the future labor market conditions, even after controlling for current economic conditions, as reflected by the gross domestic product and the level of unemployment at the time of the survey.

We first regressed the (logarithm of the) share of respondents that expect their personal job situation to worsen the next year over the (logarithm of the) unemployment rate in the month of the survey, including also country and time fixed effects.[50] This implies that the coefficients are identified only out of the variability over time within each country, and they are not influenced by time-varying factors that uniformly influence expectations across European countries.[51]

The results are reported in the first data column of Table 2, and they suggest that a 1 percent increase in the unemployment rate is associated with a 0.53 percent increase

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[49]Notable increases in the share of respondents concerned about their future personal job situation are recorded also for Cyprus (25 percentage points), Hungary (11), Ireland (14), Italy (11), Portugal (16) and Spain (12).

[50]The choice of the functional form of the equation has been informed by the choice with respect to the specification of the equation that describes the determinants of bilateral migration rates to Germany, presented in Section 5; notice that we do *not* have access to individual-level data, which would have justified to the adoption of alternative econometric models, such as an ordered probit.

[51]The adjusted R^2 of the regression of the dependent variable on the country and time fixed effects stands at 0.802.

in the share of respondents who expect a worsening of their personal job situation in the year to come. The second specification adds the (logarithm of the) interest rate on 10-year government bonds that prevails on the secondary market among the regressors: a 1 percent increase in the interest rate is associated with a 0.42 percent increase in the dependent variable, with the effect being significant at the 1 percent confidence level, while the estimated elasticity with respect to unemployment falls to 0.27. The elasticity of the expectations about the future personal job situation with respect to the interest rate is virtually unaffected when we also add the (logarithm of the) level of gross domestic product at the time of the survey to the set of regressors.

The results reported here do not depend on the specific question that we selected from the Eurobarometer survey:[52] a positive association with the yields of 10-year bonds emerges when use of the answers to *any* of the other four questions concerning the expectations for the year to come: (*i*) your life in general; (*ii*) the economic situation of your country; (*iii*) the financial situation of your household, and (*iv*) the employment situation in your country. These additional results are reported in Tables A.1-A.2 in the Appendix.

These results provide support to the hypothesis that the current interest rate on the public debt is informative about the expectations on the evolution of the economic conditions in one's own country, which might, in turn, influence the decision to migrate in a dynamic model such as the one presented in Section 2.

4   Descriptive statistics

Table 3 presents the descriptive statistics for our main sample of origin countries with respect to the rate of migration, unemployment, real GDP per capita and the 10-year bond yields. The average monthly migration rate per 1,000 inhabitants over our period of analysis stands at 0.083 throughout the sample period, with a standard deviation of 0.12. We also report an index of the migration rate, which is normalized to 100 in January 2006, to give an idea of its evolution over our seven-year period of analysis: Table 3 reveals the variability of this index, which ranges between 16.67 and 1,979.68.

The unemployment rate at origin ranges between 2.3 and 26.2 percent, and the associ-

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[52]Results are also robust to clustering the standard errors by country, and to a weighting of the estimates to reflect the differences in the sample sizes across countries.

ated index reveals that some countries have experienced a three-fold increase in the rate of unemployment since January 2006, while others reduced it by more than 50 percent. The variability in the unemployment rate is larger than the variability in quarterly real GDP per capita, with the index ranging between 81.03 and 129.89. The 10-year bond yields stand, on average, at 4.65 percent, but this average figure hides considerable variability across both time and space. Specifically, when we normalize bond yields to 100 in January 2006, we observe that the index ranges between a minimum of 24.62 and a maximum of 812.22, reflecting the diverging conditions of sovereign bond markets within the EEA in recent years.

4.1   Migration flows

Figure 2 displays gross inflows of migrants to Germany from all origin countries in the world, together with the inflows from our main sample of 28 EEA countries and with the inflows from the extended sample of 30 countries. Total gross immigration was nearly constant at around 600,000 per year between 2006 and 2009, and it then recorded a 60 percent increase up to 2012, when total inflows stood at around 965,000. Most of the observed variation in due to migration flows from EEA countries, which increased from 389,000 in 2009 to 645,000 in 2012. This implies that the countries in our main sample, which represent 62.5 percent of the inflows over our period of analysis, accounted for around 80 percent of the surge between 2009 and 2012. The main country of origin is represented by Poland (976,341 migrants over the period), followed by Romania (453,719) and Bulgaria (229,202). Some of the countries that have been more severely hit by the crisis have been climbing up the list of the main countries of origin, with Italy ranking fifth (174,592 migrants), Greece sixth (103,649) and Spain seventh (102,166).[53]

Our administrative data source also allows us to have an (imperfect) idea of the evolution over time of the stock of immigrants from each origin, a variable that is usually relied upon in the literature as a proxy for the effect of migration networks (Beine *et al.*, 2011). Specifically, we can infer the monthly evolution of bilateral migration stocks by relying on the inflows and outflows data from Statistisches Bundesamt (2013); this gives us an upper bound of the actual evolution of the stock as we do not account for the reduction in stocks due to deaths and the

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[53]For instance, although the total inflows from Greece are just 10.6 percent of the inflows from Poland over the period, Polish migration to Germany increased by 24,624 migrants between 2006 and 2012, while the corresponding increase in Greek migration stands at 25,920.

administrative data are likely to under record outflows. Migration stocks remained relatively stable between 2006 and 2012 for origin countries with a long-established migration history to Germany, such as Italy and Greece.[54] Bilateral migration stocks grew more rapidly for other origin countries, such as Poland (232,351), Romania (152,354). Bulgaria (92,719), Hungary (68,373) and Spain (34,643), as the difference between inflows and outflows amounted to 23 to 40 percent of total inflows.

5   Estimates

In this section, we present the estimates for several specifications of equation (11), where the dependent variable y_{jkt} is given by the logarithm of the ratio between gross bilateral monthly flows to Germany from the country j and the size of the population at origin. We show two sets of estimates for each specification. The first is consistent only under the restrictive assumptions on the sequential migration model that imply that the multilateral resistance to migration term r_{jkt} is identically equal to zero. In other words, it represents just a classical fixed effects (denoted FE) specification:

$$y_{jkt} = \boldsymbol{\alpha}_1^{FE\prime}\boldsymbol{x}_{jkt} + \boldsymbol{\alpha}_2^{FE\prime}\boldsymbol{x}_{jjt} + \alpha_{jk}^{FE}d_{jk} + \boldsymbol{\alpha}_t^{FE\prime}\boldsymbol{d}_t + \eta_{jkt}^{FE} \tag{14}$$

The second one is the unrestricted estimation of equation (13), denoted by CCE, that is consistent even when we account for the forward-looking and path-dependent nature of migration decisions, and that controls for a linear approximation of the r_{jkt} term through the inclusion of a vector of auxiliary regressors, as discussed in Section 2.5:

$$y_{jkt} = \boldsymbol{\alpha}_1^{CCE\prime}\boldsymbol{x}_{jkt} + \boldsymbol{\alpha}_2^{CCE\prime}\boldsymbol{x}_{jjt} + \alpha_{jk}^{CCE}d_{jk} + \boldsymbol{\alpha}_t^{CCE\prime}\boldsymbol{d}_t + \boldsymbol{\lambda}_{jk}{}'\widetilde{\boldsymbol{z}}_t + \eta_{jkt}^{CCE} \tag{15}$$

Our three potential variables of interest, included in the vector $\boldsymbol{x}_{jjt}$, which describes the utility of staying at home, are: the 10-year bond yield on sovereign debt,[55] the unemployment rate and the real GDP index of country j. All three variables enter the equation in logs.

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[54]The difference between the total inflows of Italian and Greek immigrants over the period and the recorded number of outflows stands at 16,659 and 8,953 respectively, less than 10 percent of total gross inflows for the two countries between 2006 and 2012.

[55]Given the inclusion of German (time) fixed effects, this is equivalent to including the spread and we use both terms interchangeably.

As outlined in Section 3.3, bonds yields are our measure of expectations on future earnings or, more generally, on the evolution of the economy of country j. The unemployment rate and the real GDP index are proxies for the current economic conditions in country j. Both affect employment opportunities and individual earnings. Since the unemployment rate and the real GDP index are highly correlated, we expect that multicollinearity may affect our estimation results.[56] We thus start with a more parsimonious specification taking 10-year bond yields and the unemployment rate as main explanatory variables, and consider them in addition to the real GDP index in a more comprehensive specification of the model.[57]

Although we are focusing on migration flows within the EEA where labor mobility is subject to few restrictions, we can expect some lag between changes in economic conditions or in expectations about future economic conditions and their effect on migration flows to Germany. In order to choose the empirically relevant number of lags, we follow Canova (2007) and select the optimal number of lags by running successive LR tests on dropping higher order lags.[58] The result is that we include four lags of these three variables in our monthly data regressions. What we report below is the long-run coefficient associated to each specification, that is, the sum of the lags for each of the variables.

We control for a very wide variety of other determinants of bilateral migration rates in our specifications. First, we include time fixed effects ($\boldsymbol{d}_t$). They will absorb any German-specific variation in the data as well as common elements across countries of origin over time. For example, the effect of current German economic conditions or German general migration policies is absorbed by our time fixed effects. Importantly, they also absorb the influence of the continuation payoff $V_{t+1}(k)$ on y_{jkt}. This, in turn, entails that we can form expectations on the sign of the bias due to dynamic multilateral resistance to migration, conditional upon the prevailing patterns of correlation in the data. Specifically, the derivative

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[56]Once we partial out the two variables removing the fixed effects included in the estimation, the correlation between the two variables stands at -0.60.

[57]There is also a second reason why we do not include the real GDP index in our first specification, as GDP data are available only at the quarterly level while we try to exploit the monthly variation in order to improve the precision of our estimates; as it can be seen in the summary statistics from Table 3, there is much fewer variability to exploit from a quarterly variable: the standard deviation of the normalized version of the variable is six to seven times lower than the standard deviation of the same normalized version of the unemployment rate.

[58]As suggested by Canova (2007), the highest number of lags that we included was $T^{1/3} \approx 4$, as $T = 84$ in our dataset; both the Akaike and Bayesian Information Criteria select the same number of lags.

in (6) tells us that an increase in the future attractiveness of an alternative destination unambiguously increases the continuation payoff from staying at origin at time t, $V_{t+1}(j)$, thus unambiguously reducing y_{jkt} once we control for $V_{t+1}(k)$.[59] If, say, an increase in the unemployment rate at origin is positively correlated with a reduction in the future attractiveness of some alternative destinations, this would determine an upward bias in the estimated coefficient for unemployment rate in (14), where we do not control for multilateral resistance to migration.

Second, we include origin-specific fixed effects (d_{jk}) in (14) and (15). These are introduced to control for time-invariant bilateral determinants of migration flows to Germany from a given origin. Some examples are cultural, linguistic and geographical distance, common membership in institutions that did not change over the period, and so on.

Third, our origin-specific fixed effects (d_{jk}) also partly control for slowly moving bilateral or origin-specific variables, such as the demographic composition of the population at origin. Dyadic fixed effects can also partly control for some time-varying bilateral variables, such as migration networks, for the origin countries, such as Italy, Greece and Portugal, for which these remained relatively stable over our seven-year period of analysis, as discussed in Section 4.1. As our data only provide us with an imperfect measure of the evolution of bilateral migration stocks, we follow Bertoli and Fernández-Huertas Moraga (2013) by controlling for networks through the inclusion of interactions between the dyadic fixed effects d_{jk} and dummies for sub-periods of our seven years in one of our robustness checks.[60]

Fourth, we control for two major changes in bilateral migration policies that took place during the sample period ($\boldsymbol{x}_{jkt}$). First, Romania and Bulgaria joined the EU on January 1, 2007. Although Germany did not immediately grant the free movement of workers to citizens from Bulgaria and Romania, EU membership opens numerous channels to access Germany: the freedom of settlement enables individuals to move to Germany as self-employed or small business owners. Moreover, the opportunities for seasonal work, contract work and the posting of workers have been extended in the context of EU enlargement. Finally, Germany granted persons with a university degree from the new EU member states access to its labour market. Thus, accession to the EU facilitated immigration from Bulgaria and Romania considerably albeit Germany decided to postpone the full application of the rules for the

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[59]The ambiguity would remain if we did not control for $V_{t+1}(k)$, as shown in (7).

[60]Specifically, we will introduce interactions between d_{jk} and dummies for halves (3 years and 6 months) and fourths (1 year and 9 months) of our sample.

free movement of workers until January 1, 2014. The second major policy change was the introduction of the free movement of workers in May 1, 2011 for the eight Central and Eastern European member states which joined the EU in 2004. Germany did not only introduce the free movement of workers on May 1, 2011, but abrogated also the remaining restrictions for service trade including the posting of workers, which may have further facilitated immigration from the eight Central and Eastern European member states. Immigration conditions for third-country nationals, i.e. citizens of non-EU and non-EEA countries, have remained by and large unchanged during the sample period. The new German immigration law became effective in 2005, and the 2009 amendment of this law involved only some incremental changes in channels that have been quantitatively negligible.

Notice again that our origin specific fixed effects (d_{jk}) control for migration determinants such as visa policies which are time-invariant during the sample period, while our time fixed effects ($\boldsymbol{d}_t$) control for general German migration policies that are not origin-specific.

Finally, we also control for country-specific seasonal effects in the data in our monthly specifications (also included in $\boldsymbol{x}_{jkt}$). The monthly flows we study present obvious seasonal patterns but that vary across origin countries. While the inclusion of these controls does not affect our results, they improve the fit of the models that we present. Thus, we add origin-country times month-of-the-year fixed effects to absorb these origin-specific seasonal patterns.

The rich structure of fixed effects that we rely upon for the estimation of both (14) and (15) implies that the identifying variation comes from the correlation between the origin-specific evolution of the seasonally adjusted dependent variable and of the seasonally adjusted regressors, net of common time effects.[61]

5.1 Main specifications

As a first step, we present what we can term as a classical fixed effects specification in which the log of the migration rate from a given European country to Germany is regressed on our set of controls plus a variable that proxies for current economic conditions in that country: its unemployment rate. In all specifications, we weight observations by the population of

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[61]The inclusion of interactions between dyadic dummies and dummies for time sub-periods in some specifications further reduces the identifying variation, removing the variability across sub-periods from each origin.

the origin country.[62] The results are shown in Column (1) of Table 4. The interpretation of the long-run coefficient is straightforward: a 1 percent increase in the unemployment rate at origin is associated with a 0.73 percent increase in the emigration rate to Germany.[63] The policy variables are also remarkable. The 2011 free mobility extension was associated with an increase in the migration rate to Germany of 17 percent whereas the migration rate from Romania to Bulgaria after 2007 more than quadrupled (312 percent increase).

However, our theoretical model suggests that the estimation of these coefficients is likely to be biased by the existence of multilateral resistance to migration. The dynamic nature of the decision to migrate implies that the continuation payoffs $V_{t+1}(k)$ and $V_{t+1}(j)$, contained in r_{jkt} in equation (8), matter for the decision to migrate from an origin j to a destination k at time t. Given that our dataset only includes one destination, Germany, the continuation payoff from choosing Germany is controlled for through our time dummies $\boldsymbol{d}_t$ in equation (14), while this is not the case for the continuation payoff $V_{t+1}(j)$ from choosing the origin country.

Thus, as long as the unemployment rate at origin is correlated with the future evolution of the unemployment rate in the origin country itself and in any of all possible destinations, the coefficient $\widehat{\alpha}^{FE}$ will be upward biased.[64] As we discussed in Section 2.3, the bias due to dynamic multilateral resistance to migration would disappear only if we were willing to introduce the (admittedly implausible) assumption that either individuals take myopic migration decisions or that we live in a world with no migration costs.

We add the vector of auxiliary regressors $\widetilde{\boldsymbol{z}}_t$ to the specification presented in Column (2) of Table 4 in order to remove the bias that arises in the presence of multilateral resistance to migration both of the static and of the dynamic nature.[65] The value of the F-test on

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[62]This is the theoretically relevant specification since it ensures that the location decision of each individual is weighted in the same way, independently of her country of origin. Bertoli and Fernández-Huertas Moraga (2013) or Hanson and McIntosh (2012), among others, follow this approach.

[63]Our estimate is *not* directly comparable with the one in Beine *et al.* (2013), which has a larger magnitude, as they model location-specific utility as depending on the logarithm of the employment (as opposed to the unemployment) rate, and their estimation strategy does not allow to recover the parameters of the underlying RUM model (Bertoli and Fernández-Huertas Moraga, 2012).

[64]Notice that it would also be incorrect, as discussed in Section 2.1, to interpret this coefficient as identifying the influence on the bilateral migration rate at time t of a variation in the present value of the expected stream of instantaneous utility at origin, as the individuals who decide to stay at origin at time t still retain the option to migrate at a later time.

[65]For this specification, this requires to estimate 196 additional coefficients, corresponding to 28 origin-

the assumption that the CCE term is zero is 33.71, thus strongly rejecting the null that reflects the hypothesis that r_{jkt} is identically zero for all observations. The rejection of the null confirms the need to control for multilateral resistance to migration,[66] thus suggesting that the estimated elasticity for the unemployment rate in Column (1) is biased. This is what the comparison of the estimates in Column (1) and (2) actually reveals: the long-run elasticities of migration rates to Germany with respect to origin-country unemployment rates go down from 0.73 to 0.52, a 29 percent reduction that signals that the FE estimate was upward biased. The direction of the bias is, in turn, consistent with a prevailing pattern of positive correlation in the evolution of the unemployment rate across the countries in the EEA, which could generate a static multilateral resistance to migration bias and, over time, also a dynamic multilateral resistance to migration bias. Column (2) provides a consistent estimate of the effects of unemployment rates at origin on migration rates to Germany[67] even if we have omitted relevant observable or unobservable characteristics. Unfortunately, although we include the policy variables as well, we do not have enough variation in the data as to distinguish their coefficients from those of the auxiliary CCE regressors,[68] so we cannot test our expectation on them.

The estimates in Column (2) are already free from the threat to identification posed by the dynamic multilateral resistance to migration, but it is also interesting *per se* to gain an understanding of the role played by expectations in the decision to migration to Germany, rather than just treating them as a nuisance. Thus, we add (various lags of) the yields on sovereign bonds with a residual maturity of 10 years to our vector of regressors. The results,

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specific coefficients for each of the seven auxiliary regressors.

[66]All of our results are robust to the exclusion of the policy variables from all specifications. In particular, the F-test still clearly rejects the null if policy variables are excluded or when cross-sectional averages of the policy variables are not included among the auxiliary regressors. Results available from the authors upon request.

[67]While we cannot purely establish causality, the endogeneity generated by reverse causality between changes in migration rates at the monthly level and changes in unemployment is likely to be irrelevant for most countries in the sample, given that migration rates to Germany do not generally affect great shares of the origin population.

[68]This is due to the fact that all changes in each of the two variables happen simultaneously, in January 2007 and in May 2011, respectively. This, in turn, implies that the variables for all the countries for which we have a policy change are perfectly collinear with the associated auxiliary regressors. Bertoli and Fernández-Huertas Moraga (2013) are able to identify the effect of policy changes with the CCE estimator because they have variations in the visa variable at different points in time.

corresponding to the FE and CCE estimation, are presented in Columns (3)-(4) of Table 4.

The estimates in the third data column reveal that a 10 percent increase in bond yields at origin is associated with a 3.9 percent increase in the bilateral migration rate to Germany. Interestingly, although the inclusion of our proxy for expectations should not be able to fully remove the bias due to multilateral resistance to migration, the estimated coefficient for the unemployment rate stands at 0.53, in line with the CCE estimates in Column (2) and significantly below the (biased) coefficient in Column (1).[69] This allows us to speculate that, in this particular dataset, the bias induced by dynamic multilateral resistance to migration might be more relevant than that induced by static multilateral resistance to migration.

The estimate of the coefficient on the 10-year bond yields in Column (3) is biased, as the FE specification, which does *not* fully control for multilateral resistance to migration, confounds the direct effect of a worsening in expectations at origin with the indirect effects due to a simultaneous worsening in the expectations for some alternative destinations. For instance, a positive correlation between the evolution of the bond yields for Italy and Spain would induce an upward bias in the estimated coefficient for the 10-year bond yields, as the worsening in the expectations for Spain reduces the continuation payoff $V_{t+1}(j)$ for stayers in Italy, thus indirectly increasing their incentives to move to Germany at time t.

Once we resort to the CCE estimation, Column (4) reveals that a 10 percent increase in the bond yields still significantly increases the migration rate to Germany by 1.4 percent. Recall that what we are obtaining here is a consistent estimate of what we consider the *main* component from the continuation payoff associated with remaining in the origin country: $V_{t+1}(j)$. All other elements in $V_{t+1}(j)$, corresponding to expectations in alternative destinations, are controlled for by the auxiliary regressors, which are jointly significant. The estimated elasticity of the bilateral migration rate to Germany with respect to unemployment at origin is still in line with the that we obtained from the more parsimonious specification reported in Column (2).

The last two columns in Table 4 add an additional variable (with its four lags) to the previous specifications: the log of the GDP per capita in the origin countries. The results are shown in Columns (5) and (6) of Table 4. The estimated elasticities for the 10-year bond yields variable is basically unaffected in both specifications compared to Columns (5) and

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[69]This interpretation is further corroborated by the reduction of the F-test for the specification in Column (4), which stands at 16.60, compared to Column (2).

(6). However, there is some evidence that both the unemployment rate and the GDP per capita at origin are proxying for the same concept of current economic conditions at origin. While both variables are significant in the FE specification in Column (5), the coefficient on the unemployment rate variable is significantly reduced with respect to specification (3): from 0.53 to 0.40. The estimated elasticity of the GDP per capita variable stands at -0.83 in Column (6), and it is significant at the 5 percent confidence level. As happened with the unemployment rate, multilateral resistance to migration unduly magnifies its estimated effect in the traditional FE specification reported in Column (5), where the estimated elasticity stands at -1.44.

5.2 Robustness checks

We present some relevant robustness checks on our results, where we (i) extend the sample size to 30 countries, (ii) do not weight observations, (iii) we drop Estonia as bond yields have been estimated for this country, (iv) we rely on alternative measures of population at origin, and (v) we more fully control for the possible confounding effects of networks through a richer structure of fixed effects. For all points (i) to (v) we report the FE and the CCE estimates including the rate of unemployment at origin, 10-year bond yields and immigration policy variables.

First, we test whether our results are robust if we expand our sample by adding Turkey and Croatia in order to achieve the minimum number of countries suggested by Pesaran (2006). As mentioned above, we add Turkey because it is one of the main traditional origins for German immigration although it is only the fifth country by the size of total gross inflows over our seven-year period (189,314 immigrants to Germany). Croatia is also a natural choice for our extended sample, as it joined the EU, and hence the EEA, after the end of our period of analysis on July 1, 2013, and it recorded 67,722 immigrants to Germany between 2006 and 2012. The results on this expanded 30-country sample are shown in Table 5. Columns (1) and (2) present the FE and CCE specification, which are basically unchanged with respect to our baseline specifications in Columns (3) and (4) in Table 4. The estimated elasticity for the 10-year bond yields is slightly larger than in our baseline, but the difference is not significant at conventional confidence levels.

We reran the model on the baseline sample without population weights in Columns (3) and (4) in Table 5. The estimated coefficients are again barely different from our baseline

in Table 4. Then, we also test what happens when we estimate both the FE and CCE specifications dropping Estonia, the only country in our main sample for which we did not have actual bond yields data and had to estimate them. As it could be expected, the omission of a country that represents barely more than 0.3 percent of the total population of the included origin countries has no impact whatsoever on the estimates. Columns (5) and (6) from Table 5 are almost identical to specifications (3) and (4) from Table 4.

The first two specifications in Table 6 report the results that we obtain when using population figures from EUROSTAT (2013c) to relate observed gross migration flows to Germany to the size of the population at origin aged 15 to 49, which probably represents the age range from which most recent migrants to Germany are drawn from. Columns (1) and (2) reveal that this change in the definition of the dependent variable has no impact on our estimates. In unreported results, we have also experimented with alternative definitions of the denominator of the migration rate closer to the theoretical concept of those remaining in the origin countries and interpolating monthly observations so as to obtain the same frequency as in our migration data. The results remain virtually unchanged due to the fact that most of the variation in the migration rate comes from the numerator independently of the definition of the denominator.[70]

The other specifications in Table 6 include the interaction between the origin dummies d_{jk} and dummies for sub-periods of our sample, whose length is respectively 42 and 21 months. This greatly reduces the variability that we use for identification, allowing us to better control for the effects of time-varying migration networks at destination and other potential time-varying confounders. Our estimates are robust to this richer structure of fixed effects, whose most relevant effect is to reduce the size of the bias in the estimated coefficient for bond yields in the FE specification.[71]

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[70]Results available from the authors upon request.

[71]The magnitude of the estimated effect of the 2007 EU enlargement is reduced in size, as we use a more limited number of months after the policy change to identify its effects; this reveals that the effects of the policy change might have emerged gradually, although the FE estimates are likely to be still biased because of multilateral resistance to migration.

6   Concluding remarks

This paper has analyzed the effects of the economic crisis that began in 2008 on migration flows from EEA countries to Germany, tackling the econometric challenges that arise from the sequential model of migration that we developed from the seminal contribution of Kennan and Walker (2011). We control for the confounding influence exerted by the future attractiveness of alternative destination countries with the econometric approach proposed by Bertoli and Fernández-Huertas Moraga (2013), and we extend the usual definition of location-specific utility to better account for the role of expectations in shaping the decision to migrate.

We find evidence that economic conditions at origin and the expectations about their evolution, proxied by the yields on 10-year government bonds, significantly influence the scale of bilateral migration flows. The estimation of both effects is biased when we resort to more restrictive estimation approaches, which are *not* consistent with the forward-looking and path-dependent character of the decision to migrate, and the direction of the bias is consistent with the predictions generated by the underlying RUM model of migration. This confirms that variations in the future attractiveness of alternative destinations do play a role in shaping current migration decisions, thus suggesting that, say, the crisis in Spain contributes to explain not only the increase in Spanish migration, but also the surge in Romanian and Bulgarian migration to Germany observed since 2006. This, in turn, implies that the economic crisis played an important and subtle influence in reshaping the landscape of international migration flows within Europe that goes beyond its direct effects on the citizens of the countries that have been more severely affected.

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Table 1: Descriptive statistics from the Eurobarometer surveys

Country	*worse personal job situation (percent)*				
	Average	St. dev.	Lowest	Highest	Change
Austria	8.3	1.6	6	11	-3
Belgium	7.5	2.2	5	11	2
Bulgaria	11.5	2.7	8	17	-7
Cyprus	12.6	6.5	7	33	25
Czech Republic	12.4	3.0	8	18	8
Denmark	2.9	0.8	2	5	-1
Estonia	7.1	2.8	3	12	0
Finland	3.7	1.3	2	6	2
France	6.2	1.4	4	10	-1
Germany	7.5	2.2	4	12	-7
Greece	21.7	13.0	8	42	27
Hungary	26.7	6.0	16	37	11
Iceland	3.5	1.6	2	6	-
Ireland	12.0	5.2	4	20	14
Italy	15.9	3.7	11	22	11
Latvia	9.3	5.8	4	25	0
Lithuania	21.6	9.1	11	40	-4
Luxembourg	5.0	2.0	4	11	2
Netherlands	7.7	1.9	6	12	5
Poland	9.9	3.4	4	16	2
Portugal	23.4	7.1	14	38	16
Romania	14.9	5.6	8	26	5
Slovak Republic	13.7	3.4	9	20	3
Slovenia	11.1	3.4	6	18	9
Spain	10.2	4.1	5	18	12
Sweden	3.5	0.8	2	4	-1
United Kingdom	6.9	2.1	5	11	0

Notes: values are in percent; change reports the difference in percentage points between the November 2012 and the April 2006 surveys; 14 observations for each country, except for Iceland (6 observations), as Iceland has been covered by the Eurobarometer only since 2010.
Source: Authors' elaboration on data from European Commission (various issues).

Table 2: Determinants of the expectations on your personal job situation

	Dependent variable:		
	worse personal job situation		
	Specification		
Variables	(1)	(2)	(3)
Unemployment rate	0.526	0.267	0.193
	[0.070]***	[0.074]***	[0.090]**
10-year bond yields		0.415	0.397
		[0.056]***	[0.057]***
Gross domestic product			-0.704
			[0.489]
Adjusted R^2	0.831	0.856	0.856
Observations	330	330	330
Countries	26	26	26
Surveys	14	14	14
Country dummies	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$; standard errors in brackets; the dependent variable is the share of respondents who expect their personal job situation to worsen over the next year; all variables are in natural logarithms.
Source: Authors' elaboration on data from European Commission (various issues) and the data described in Section 3.2.2.

Table 3: Descriptive statistics

Variable	Average	St. dev.	Min	Max	Obs.
Migration rate ($\times$ 1,000)	0.083	0.12	0.001	0.89	2,352
Migration rate index	167.54	131.05	16.67	1,979.68	2,352
Unemployment rate	8.88	4.16	2.3	26.2	2,352
Unemployment rate index	113.64	49.31	42.24	343.18	2,352
Real GDP index	103.79	6.40	81.03	129.89	2,352
10-year bond yields	4.65	1.95	0.61	29.24	2,352
10-year bond yield index	120.23	51.64	24.62	812.22	2,352

Notes: monthly series on 28 countries over the period January 2006 - December 2012; all descriptive statistics are weighted by population at origin; all indices are set to 100 in January 2006.

Source: Authors' elaboration on Statistisches Bundesamt (2013) and on the data presented in Section 3.2.2.

Table 4: Determinants of bilateral migration rates to Germany (2006-2012)

			Dependent variable:			
			log of the bilateral migration rate to Germany			
			Specification			
Model	FE	CCE	FE	CCE	FE	CCE
Variables	(1)	(2)	(3)	(4)	(5)	(6)
log unemployment rate	0.729	0.521	0.533	0.486	0.400	0.384
(4 lags)	[0.021]***	[0.035]***	[0.019]***	[0.043]***	[0.022]***	[0.059]***
log 10-year bond yields			0.390	0.141	0.350	0.142
(4 lags)			[0.015]***	[0.028]***	[0.015]***	[0.032]***
log GDP per capita					-1.443	-0.830
(4 lags)					[0.125]***	[0.350]**
2007 EU Enlargement	1.138	(included)	1.053	(included)	1.135	(included)
	[0.038]***		[0.033]***		[0.033]***	
2011 Free Mobility	0.172	(included)	0.135	(included)	0.231	(included)
	[0.021]***		[0.018]***		[0.020]***	
CCE Test (*p*-value)		33.71 (0.00)		16.60 (0.00)		12.01 (0.00)
Time dummies	yes	yes	yes	yes	yes	yes
Origin-month dummies	yes	yes	yes	yes	yes	yes
Auxiliary regressors	no	yes	no	yes	no	yes
Adjusted R^2	0.980	0.995	0.985	0.995	0.986	0.995
Countries	28	28	28	28	28	28
Observations	2,352	2,352	2,352	2,352	2,352	2,352

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$; standard errors in brackets; observations weighted by population at origin; the coefficients on the independent variables are the long-run coefficients on four lags of the variables; CCE specifications include the cross-sectional averages of the dependent and independent variables interacted with origin fixed effects as auxiliary regressors; policy variables are included but not identified because of perfect collinearity with the auxiliary regressors in the CCE specifications; the null of the CCE test (*F*-test) is that the origin-specific coefficients of the cross-sectional averages included as auxiliary regressors are jointly zero.

Source: Authors' elaboration on Statistisches Bundesamt (2013) and on the data presented in Section 3.2.2.

Table 5: Determinants of bilateral migration rates to Germany, robustness checks

	Dependent variable:					
	log of the bilateral migration rate to Germany					
	Specification					
	Adds Croatia and Turkey		*No population weights*		*Drops Estonia*	
Model	FE	CCE	FE	CCE	FE	CCE
Variables	(1)	(2)	(3)	(4)	(5)	(6)
log unemployment rate	0.539	0.492	0.491	0.373	0.533	0.483
(4 lags)	[0.018]***	[0.039]***	[0.024]***	[0.072]***	[0.019]***	[0.044]***
log 10-year bond yields	0.394	0.198	0.330	0.162	0.391	0.141
(4 lags)	[0.015]***	[0.025]***	[0.020]***	[0.044]***	[0.015]***	[0.029]***
2007 EU Enlargement	1.090	(included)	1.177	(included)	1.054	(included)
	[0.034]***		[0.049]***		[0.033]***	
2011 Free Mobility	0.157	(included)	0.362	(included)	0.134	(included)
	[0.019]***		[0.023]***		[0.019]***	
CCE Test (p-value)		14.81 (0.00)		9.76 (0.00)		16.82 (0.00)
Time dummies	yes	yes	yes	yes	yes	yes
Origin-month dummies	yes	yes	yes	yes	yes	yes
Auxiliary regressors	no	yes	no	yes	no	yes
Weights	yes	yes	no	no	yes	yes
Adjusted R^2	0.983	0.994	0.964	0.984	0.985	0.995
Countries	30	30	28	28	27	27
Observations	2,520	2,520	2,352	2,352	2,268	2,268

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$; standard errors in brackets; observations weighted by population at origin when noted; the coefficients on the independent variables are the long-run coefficients on four lags of the variables; CCE specifications include the cross-sectional averages of the dependent and independent variables interacted with origin fixed effects as auxiliary regressors; policy variables are included but not identified because of perfect collinearity with the auxiliary regressors in the CCE specifications; the null of the CCE test (F-test) is that the origin-specific coefficients of the cross-sectional averages included as auxiliary regressors are jointly zero.

Source: Authors' elaboration on Statistisches Bundesamt (2013) and on the data presented in Section 3.2.2.

Table 6: Determinants of bilateral migration rates to Germany, robustness checks (cont'd)

	Dependent variable:					
	log of the bilateral migration rate to Germany					
	Specification					
	Population aged 15-49		*Two sub-periods*		*Four sub-periods*	
Model	FE	CCE	FE	CCE	FE	CCE
Variables	(1)	(2)	(3)	(4)	(5)	(6)
log unemployment rate	0.522	0.497	0.465	0.490	0.396	0.462
(4 lags)	[0.019]***	[0.043]***	[0.021]***	[0.046]***	[0.028]***	[0.055]***
log 10-year bond yields	0.404	0.138	0.310	0.147	0.213	0.137
(4 lags)	[0.015]***	[0.029]***	[0.015]***	[0.030]***	[0.019]***	[0.031]***
2007 EU Enlargement	1.043	(included)	0.821	(included)	0.730	(included)
	[0.032]***		[0.028]***		[0.030]***	
2011 Free Mobility	0.131	(included)	0.205	(included)	0.172	(included)
	[0.018]***		[0.017]***		[0.049]***	
CCE Test (p-value)		17.17 (0.00)		6.66 (0.00)		2.82 (0.00)
Time dummies	yes	yes	yes	yes	yes	yes
Origin-month dummies	yes	yes	yes	yes	yes	yes
Origin-sub-period dummies	no	no	yes	yes	yes	yes
Auxiliary regressors	no	yes	no	yes	no	yes
Adjusted R^2	0.985	0.995	0.991	0.995	0.994	0.995
Countries	28	28	28	28	28	28
Observations	2,352	2,352	2,352	2,352	2,352	2,352

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$; standard errors in brackets; observations weighted by population at origin; the coefficients on the independent variables are the long-run coefficients on four lags of the variables; CCE specifications include the cross-sectional averages of the dependent and independent variables interacted with origin fixed effects as auxiliary regressors; policy variables are included but not identified because of perfect collinearity with the auxiliary regressors in the CCE specifications; the null of the CCE test (F-test) is that the origin-specific coefficients of the cross-sectional averages included as auxiliary regressors are jointly zero.

Source: Authors' elaboration on Statistisches Bundesamt (2013) and on the data presented in Section 3.2.2.

Appendix

A Further results from the Eurobarometer surveys

Table A.1: Determinants of expectations

	Dependent variable:					
	worse life in general			worse economic situation in the country		
	Specification					
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Unemployment rate	0.473	0.340	0.328	0.231	0.129	0.145
	[0.072]***	[0.080]***	[0.097]***	[0.077]***	[0.087]	[0.106]
10-year bond yields		0.214	0.2171		0.164	0.167
		[0.061]***	[0.062]***		[0.066]**	[0.068]**
Gross domestic product			-0.115			0.153
			[0.528]			[0.575]
Adjusted R^2	0.849	0.854	0.854	0.606	0.612	0.611
Observations	356	356	356	356	356	356
Countries	26	26	26	26	26	26
Surveys	14	14	14	14	14	14
Country dummies	Yes	Yes	Yes	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes	Yes

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$; standard errors in brackets; the dependent variable in specifications (1)-(3) is the share of respondents who expect their life in general to worsen over the next year, and in specifications (4)-(6) the share of respondents who expect the economic situation in their country to worsen over the next year; all variables are in natural logarithms.

Source: Authors' elaboration on data from European Commission (various issues) and the data described in Section 3.2.2.

Table A.2: Determinants of expectations (cont'd)

	Dependent variable:					
	worse household financial situation			worse employment situation in the country		
	Specification					
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Unemployment rate	0.452	0.222	0.193	0.463	0.414	0.480
	[0.065]***	[0.069]***	[0.084]***	[0.091]***	[0.095]***	[0.116]***
10-year bond yields		0.370	0.363		0.080	0.096
		[0.052]***	[0.054]***		[0.072]	[0.074]
Gross domestic product			-0.272			0.632
			[0.455]			[0.629]
Adjusted R^2	0.799	0.826	0.826	0.634	0.635	0.619
Observations	356	356	356	356	356	356
Countries	26	26	26	26	26	26
Surveys	14	14	14	14	14	14
Country dummies	Yes	Yes	Yes	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes	Yes

Notes: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$; standard errors in brackets; the dependent variable in specifications (1)-(3) is the share of respondents who expect their the financial situation of their household to worsen over the next year, and in specifications (4)-(6) the share of respondents who expect the employment situation in their country to worsen over the next year; all variables are in natural logarithms.
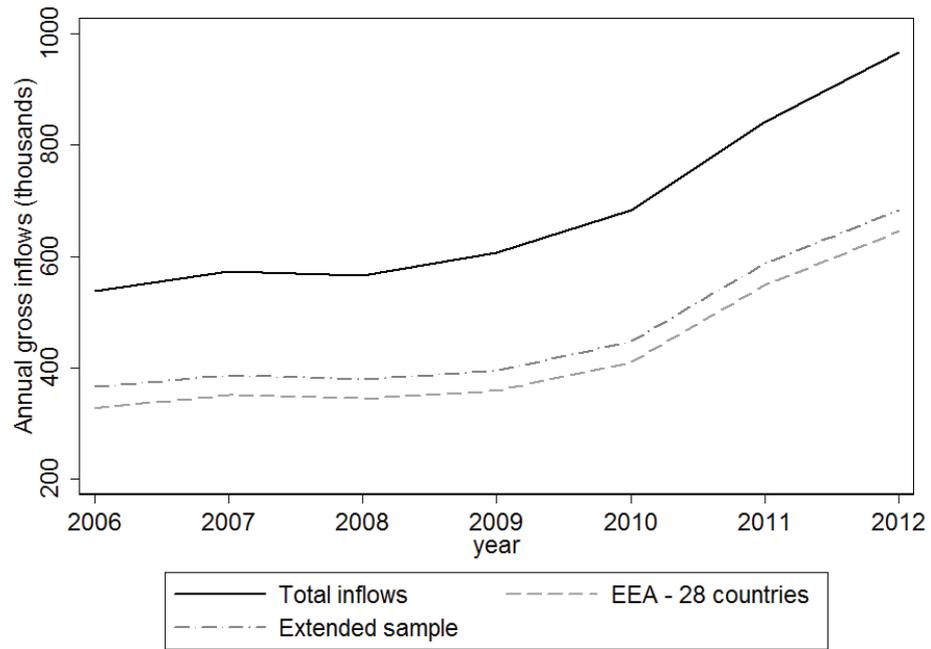
Source: Authors' elaboration on data from European Commission (various issues) and the data described in Section 3.2.2.

Figure 1: Gross flows from Romania to Germany and Spain, 2006-2012
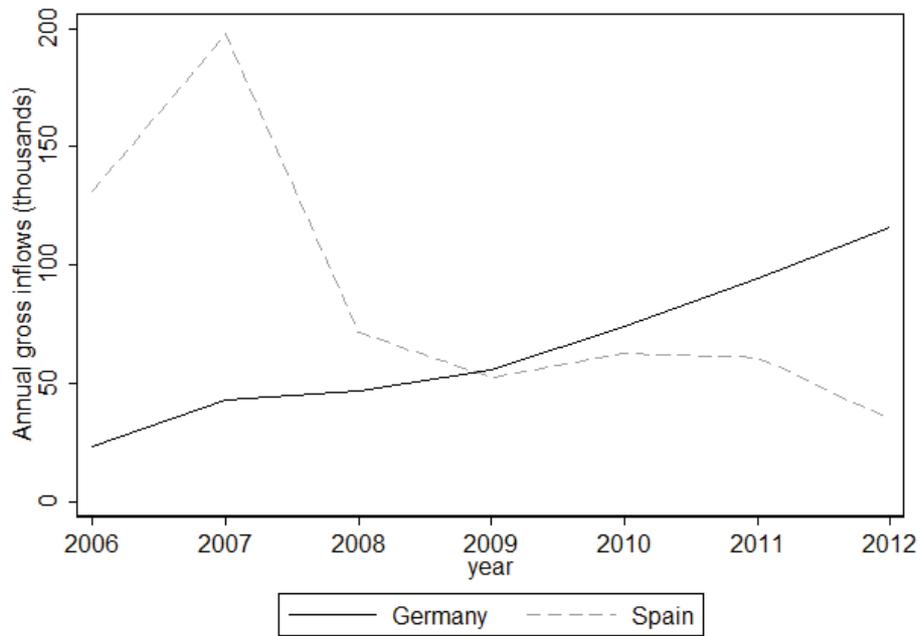
Figure 2: Gross inflows to Germany by country sample, 2006-2016

Source: Authors' elaboration on Statistisches Bundesamt (2013).