

THE IMPACT OF LABOR MARKET DYNAMICS ON THE RETURN MIGRATION OF IMMIGRANTS

Govert E. Bijwaard, Christian Schluter, and Jackline Wahba*

Abstract—Using administrative panel data on the entire population of new labor immigrants to the Netherlands, we estimate the effects of individual labor market spells on immigration durations using the timing-of-events method. The model allows for correlated unobserved heterogeneity across migration, unemployment, and employment processes. We find that unemployment spells increase return probabilities for all immigrant groups, while reemployment spells typically delay returns.

I. Introduction

THE labor market performance of immigrants in the host country has received ample attention in the empirical literature. Neglected, however, is the question as to what extent this labor market performance affects the decisions of migrants to return to their source country. In particular, what is the effect of adverse or positive labor market events such as the occurrence of unemployment spells and reemployment spells? How does the effect vary by the duration of the labor market spell? The failure of the empirical literature to ask these questions, let alone to furnish convincing answers, arises from a combination of methodological challenges and severe limitations of the data usually encountered in migration analysis.

We address these novel questions using a unique administrative panel for the entire population of recent immigrants to the Netherlands covering the years 1999 to 2007. These data characteristics—large size, repeated and accurate measurement—are fairly unique in migration analysis (exceptions are Nekby, 2006, for Sweden and Aydemir and Robinson, 2008, for Canada) and enable us to examine durations reliably. The usual data situation is one of small samples, possibly subject to selectivity and attrition issues, extracted from surveys of respondents who provide recall data. These problems are particularly acute in studies of migration durations since survey attrition usually confounds out-migration. By contrast, our administrative population data have no attrition. We expand on three important features of our data.

First, this Dutch immigrant register is based on the legal requirement for immigrants to register with the authorities

upon arrival.¹ Several other official registers are linked by Statistics Netherlands to this immigrant register, such as social benefits and the income register (used by the tax authorities). Sojourn times in labor market states are thus accurately recorded. Consequently, no data based on individual recall have to be used. Moreover, the usual concerns about measurement error are less acute.

A second important feature of our data is the administrative report in the immigrant register (consistent with the visa status at entry) of the reason for immigration. This enables us to focus explicitly and exclusively on 94,270 labor immigrants. The reason for the immigration is usually latent in standard data sets, and different behavioral patterns of labor and nonlabor migrants would confound the empirical analysis. Indeed, the other immigrant groups, such as family migrants, differ systematically from labor migrants in terms of labor market attachment, return behavior, and demographic characteristics.

Size is the third attractive feature of our data to be highlighted here, as it allows us to stratify our analysis rather than be constrained to estimate one common model on pooled data. Such pooling is problematic since immigration laws stipulate visa requirements that differ by country of origin, and visa status in turn affects migration and labor market behaviors. We consider immigrants from sending countries in the EU15 (“old Europe”) and the new EU (the majority of whom arrived after the EU enlargement in 2004); the countries outside Europe are grouped into developed (DCs) and less developed (LDCs) sending countries.

While modern duration analysis (see van den Berg, 2001, for a survey) is widely applied in labor economics, the limitations of survey data have prevented its widescale adoption in migration studies. The richness of our data enables us to go beyond standard modeling of migration durations and to tackle the complex task of examining jointly the migration and labor market processes using a mixed proportional hazards framework. In particular, we estimate the effects of the labor market dynamics on the return decision of immigrants using the “timing-of-events” method (Abbring & van den Berg, 2003). At the same time, we control for the correlated effects that arise from the correlation between unobservables in the migration and labor market process (ignoring these, we show, would result in substantial endogeneity biases). Our model for the migration duration permits the sojourn times in the various labor market states to have effects that depend on both duration and the timing of the labor market spells.

¹ They are also required to deregister when leaving. Noncompliers are removed from the register by administrators. We know that these individuals have out-migrated, but the exact date of their departure is unknown. We take this administrative removal into account in our modeling.

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* Bijwaard: Netherlands Interdisciplinary Demographic Institute; Schluter: Aix-Marseille Université (Aix Marseille School of Economics), CNRS and EHESS, and University of Southampton; Wahba: University of Southampton.

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These concerns distinguish this paper from the recent empirical literature on return migration. For instance, out-migration propensities of both immigrants and natives have been considered in Nekby (2006) for Sweden. Using administrative data, she estimates linear probability models and focuses on the role of characteristics, finding evidence of positive selection in terms of education. Rather than using a static framework, Aydemir and Robinson (2008) estimate standard proportional hazard models on administrative data for Canada, reporting evidence of positive selection in terms of skill levels. Instead of focusing exclusively on characteristics, Kirdar (2009) also considers the effect of unemployment spells on the return decision using survey data for Germany. Estimating logistic hazard specifications, he finds that the effect of unemployment spells on out-migration hazards depends on unemployment durations. Bijwaard (2010) estimates mover-stayer proportional hazards models using administrative data for different migration motives for the Netherlands, reporting high return rates for labor migrants. However, he does not consider labor market dynamics, as we do here. These approaches to migration duration estimation differ from ours. Unlike these three studies, we also control for unobserved heterogeneity (see van den Berg, 2001, for a discussion of the importance of this). A complementary, mainly theoretical literature focuses on the reason for return migration. However, by not considering the effects of labor market shocks, the concerns of this literature are different from ours.²

We find that, unconditionally, both unemployment and return migration are substantial: between 35% and 50% of labor immigrants experience unemployment, and 48% leave the host country during the observation window of 1999 to 2007. Turning to the effects of the labor market experiences, we found that unemployment periods increase return probabilities for all immigrant groups, while reemployment typically delays returns. All effects are substantial and significant.

The impact estimates are valid under the assumption that unemployment affects the migration hazard only after its start, not before. We discuss several threats to identification posed by the potential anticipation of the unemployment event by out-migration (noting that anticipators, by terminating their employment, forgo income and the option value of on-the-job search). Further threats in the Dutch case arise at precise migration durations that trigger statutory changes in employment and residence rights of the individual. The scope for systematic anticipation is, however, limited by the flexibility of the Dutch labor market, the little power that unions have, and the short statutory notice period. Finally, we verify the robustness of our results to several of our population selection

choices in several data- and simulation-based sensitivity checks.

The outline of the paper is as follows. The data are described in section II. The econometric model is set out in detail in section III. We specify the labor market and the migration processes and elucidate the role of unobservable heterogeneity. The identification argument is discussed in section IIIC. Estimation proceeds by maximizing the likelihood, presented in section IIIE. The empirical results follow in section IV, and in section IVE, we present several robustness checks. Section V concludes.

II. Administrative Panel Data on the Population of Recent Immigrants to the Netherlands

All legal immigration by non-Dutch citizens to the Netherlands is registered in the Central Register Foreigners (Centraal Register Vreemdelingen, CRV), using information from the Immigration Police (Vreemdelingen Politie) and the Immigration and Naturalization Service (Immigratie en Naturalisatie Dienst, IND). It is mandatory for every immigrant to notify the local population register immediately after arrival in the Netherlands if he or she intends to stay for at least two-thirds of the following six months. Natives as well as immigrants are required to register with their municipality. Our data comprise the entire population of immigrants who entered during our observation window of 1999 to 2007.

In addition to the date of entry and exit, the administration records the individual's reason from migration. The motive is usually coded according to the visa status of the immigrant (see the web appendix for some details); if not, the immigrant reports the reason when registering in the population register. Statistics Netherlands distinguishes several reasons: labor migrants, family migrants, student immigrants, asylum seekers (and refugees), and immigrants for other reasons. Given the focus of this paper on labor market events, we consider exclusively labor migrants, which represent 26% of all non-Dutch immigrants in the age group 18 to 64.³

It is possible that the reason for labor migration is either miscoded or misreported. Since most non-EU labor migrants require an employment-dependent work visa to immigrate, they should be formally employed shortly after entry. Thus, in order to limit the possibilities of misclassification error of the reason for labor migration, we require that immigrants be employed in the Netherlands within three months of their entry (we have verified that our results are not sensitive to this particular cut-off; see section IVE). Of the original 146,290 migrants with a labor motive, 94,270 (64%) are employed within three months of their entry.

This immigration register is linked by Statistics Netherlands to the Municipal Register of Population (Gemeentelijke

²The two principal opposing paradigms are theories of optimal migration durations based on preference for source country consumption (Galor & Stark, 1991; Dustmann & Weiss, 2007) or target savings (Dustmann, 2003), and theories of mistaken expectations and immediate failure on the labor market (Borjas & Bratsberg, 1996). Empirical tests are conducted in Yang (2006) and Gibson and McKenzie (2011).

³This restriction is justified in detail in the discussion paper version of this paper (Bijwaard, Schluter, & Wahba, 2012), where we show that the other migrant groups have a substantially larger propensity to remain in the Netherlands, have a significantly weaker attachment to the labor market, and are demographically different.

Basisadministratie, GBA) and their Social Statistical Database (SSD). The GBA contains basic demographic characteristics of the migrants, such as age, gender, marital status, and country of origin. From the SSD we have information (on a monthly basis) on the labor market position, income, industry sector, housing, and household situation. Since we consider only new entrants to the Netherlands, most immigrants are not eligible for social benefits such as unemployment insurance payments; these are conditional on sufficiently long employment or length of residence (for non-EU immigrants, essentially five years). Statistics Netherlands classifies an individual as employed if the principal source of income is employment and as unemployed otherwise.

Although in principle the exact date of emigration is known, some migrants do not officially inform the authorities when they leave. The departure of these noncomplying individuals is registered as an administrative removal after the authorities have assessed that the migrant has left the municipality without showing up in the files of another municipality in the Netherlands or as an emigrant. Given the municipality registration requirement for all citizens, we know that these migrants have left the Netherlands for certain; however, the exact date of their departure is unknown. These administrative removals are included among emigration, and they amount to around 38% of all emigrations. Of these administratively removed migrants, 73% have no observed income in the country. We conjecture that the majority of these migrants left the country shortly after they stopped receiving income (either earnings or benefits). For those who still have income until they are administratively removed, we assume that they left at that exact date. We have explicitly addressed the issue of administrative removals in the formulation of the likelihoods below.

We thus arrive at a data set that contains 94,270 labor immigrants, 120,287 employment spells, and 56,783 periods of unemployment. Distinguishing migrants according to their initial labor mobility, we estimate separate models for migrants from sending countries in the EU15 (“old Europe”), the new EU (the majority of whom are Poles and arrived after 2004), and the countries outside Europe, which are grouped into developed (DCs) and less developed (LDCs) sending countries. These groupings are defined precisely in the online appendix.

A. Summary Statistics: Labor Immigrants

We discuss the summary statistics for our data relating to the dynamics of migration and of labor market events (the online appendix considers other aspects of the data). In table 1 we consider the incidence of return migration and conditional on returning, the duration of the stay in the Netherlands. Note that the group of “stayers” includes permanent immigrants and temporary migrants who have not yet returned. Hence immigrants from the new EU, having arrived predominantly in the second half of our observation window, are expected to exhibit a high proportion of incomplete migration periods.

TABLE 1.—DESCRIPTIVE DYNAMICS

	EU15	New EU	Non-EU	
			DCs	LDCs
Stayer ^a	48.2%	70.9%	40.9%	58.6%
<i>Length of stay at return migration</i>				
Less than 6 months	3.9%	12.9%	2.4%	6.0%
6–12 months	11.3%	20.0%	10.3%	13.6%
12–18 months	13.3%	17.1%	12.4%	13.2%
18–24 months	13.1%	14.0%	15.2%	13.4%
24–60 months	46.4%	29.7%	49.1%	43.6%
More than 5 years	11.6%	6.4%	10.8%	10.2%
Average in months	32.5	23.9	32.8	30.6
<i>Labor market dynamics</i>				
<i>Mean number of periods of employment and unemployment per migrant</i>				
Employment	1.349	1.233	1.093	1.221
Unemployment	0.714	0.475	0.405	0.509
Always employed ^b	49.9%	63.8%	64.5%	62.6%
Unemployed once	36.9%	28.5%	32.0%	29.1%
Unemployed more than once	13.2%	7.8%	3.5%	8.3%
Never reemployed ^c	77.9%	83.4%	92.9%	86.2%
Reemployed once	14.5%	12.2%	5.8%	9.1%
Reemployed more than once	7.5%	4.4%	1.3%	4.8%

^aStayers are migrants who remain in the country to the end of the observation period.

^bPercentage of migrants employed through their entire stay in the country.

^cPercentage of migrants never reemployed, unemployed, and then reemployed during their stay in the country.

This is borne out in the data, since the share of stayers from the new EU is 71%, whereas for other immigrants the range is between 41% and 59%. Relatedly, the durations of their completed periods of employment are shorter. However, a large share of new EU movers (13%) leave the Netherlands after less than six months, which is considerably larger than for other immigrant groups. These differences highlight the importance of an analysis disaggregating by sending countries. Immigrants from the EU15 are more (less) likely to stay than migrants from (less) developed countries outside the EU. Conditional on returning, the distribution of completed durations looks fairly similar for these three groups, as do the average durations.

Turning to the return probabilities by migration duration, figure 1 depicts their Kaplan Meier estimates for all immigrant groups.⁴ All groups look very similar for durations up to 24 months, after which the estimates fan out. Considering the two extreme estimates, immigrants from non-EU DCs are more likely to return than new EU immigrants, and this gap increases significantly until the end of the observation window. The other immigrant groups assume an intermediate position. Overall, both table 1 and figure 1 highlight the temporary nature of labor migration. Across all immigrant groups, a substantial proportion leave the Netherlands within our observation window of 1999 to 2007, and many do so within 24 months.⁵

⁴ See Wooldridge (2010). The estimator is $Pr\{T_m \leq a_r\} = 1 - \prod_{i=1}^r (N_i - E_i)/N_i$ with $r = 1, \dots, R$ where T_m denotes the migration duration, E_i the number of immigrants observed to leave in the i th interval $[a_{i-1}, a_i)$, and N_i is size of the risk set for this interval (the number of immigrants who have neither left nor have been censored at time a_{i-1}).

⁵ As a comparison, Aydemir and Robinson (2008) find that the out-migration rate in Canada twenty years after arrival is around 35% among young, working-age, male immigrants, and around six of ten of those who leave do so within the first year of arrival.

FIGURE 1.—KAPLAN MEIER ESTIMATES OF RETURN PROBABILITIES

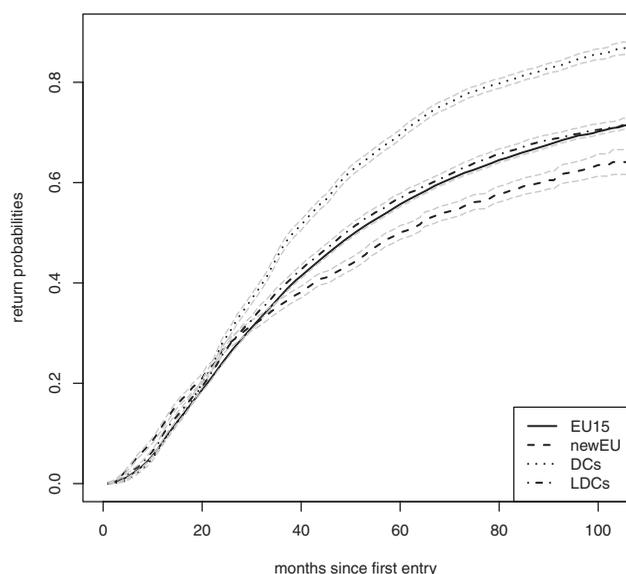


TABLE 2.—DESCRIPTIVE STATISTICS: UNEMPLOYED IMMIGRANTS WHO LEAVE

	EU15	New EU	Non-EU	
			DCs	LDCs
Unemployed at emigration ^a	54.3%	44.4%	40.8%	48.7%
Repeated unemployment ^b	20.4%	14.1%	7.9%	15.1%
Mean number & period of unemployment ^b	1.30	1.20	1.10	1.22
<i>Current unemployment duration^b</i>				
Less than 3 months	18.5%	26.4%	23.3%	20.1%
3–6 months	15.8%	22.5%	17.1%	17.1%
6–12 months	22.7%	22.6%	19.5%	22.4%
More than 1 year	43.0%	28.5%	40.1%	40.4%
Average number of months	15.1	10.8	14.6	13.8
<i>Preceding employment duration^b</i>				
Less than 3 months	16.4%	17.7%	6.2%	11.8%
3–6 months	15.1%	17.7%	10.8%	11.9%
6–12 months	45.7%	39.0%	60.4%	54.0%
More than 1 year	22.8%	25.6%	22.6%	22.3%
Average number of months	15.5	13.2	20.0	17.8

^aAs percentage of all migrants who leave.

^bAt the departure moment for the migrants who are unemployed when they leave.

Turning to the labor market dynamics, table 1 reveals that migrants from the EU15, relative to the other groups, experience greater labor market volatility. During the observation window, they experience a higher incidence of unemployment (the mean number of periods of unemployment is 0.7), and more employment periods (1.3). A slightly bigger proportion are more than unemployed once (13%), and the share of the “always employed” is smaller (50%).⁶

Since we seek to estimate the effects of negative and positive individual labor market shocks on the migration durations, we now consider the immigrants by their labor market status prior to their departure from the host country. Tables 2 and 3 condition on leaving the host country, whereas table 1 considered the unconditional labor market dynamics.

⁶Note that migrants who are always employed enter the likelihood as censored observations.

TABLE 3.—DESCRIPTIVE STATISTICS: REEMPLOYED IMMIGRANTS WHO LEAVE

	EU15	New EU	Non-EU	
			DCs	LDCs
Reemployed at emigration ^a	7.4%	9.9%	2.7%	4.8%
Repeated reemployment ^b	29.9%	29.3%	10.2%	33.8%
Mean number of periods of reemployment ^b	1.45	1.42	1.13	1.52
<i>Current (re-)employment duration^b</i>				
Less than 3 months	18.2%	26.9%	17.1%	17.0%
3–6 months	16.7%	21.6%	15.1%	17.4%
6–12 months	23.4%	22.5%	26.9%	23.7%
More than 1 year	41.7%	29.0%	40.9%	41.9%
Average number of months	15.0	10.7	13.9	14.4
<i>Preceding unemployment duration^b</i>				
Less than 1 month	18.3%	22.2%	11.3%	20.7%
1–2 months	18.1%	18.1%	18.3%	13.2%
2–3 months	17.2%	20.0%	16.7%	20.4%
3–6 months	20.5%	19.7%	17.4%	22.0%
6–12 months	17.4%	13.2%	26.3%	15.9%
More than 1 year	8.5%	6.8%	8.6%	7.8%
Average number of months	5.1	4.3	6.3	4.9

^aAs percentage of all migrants who leave.

^bAt the departure moment for the migrants who are reemployed when they leave.

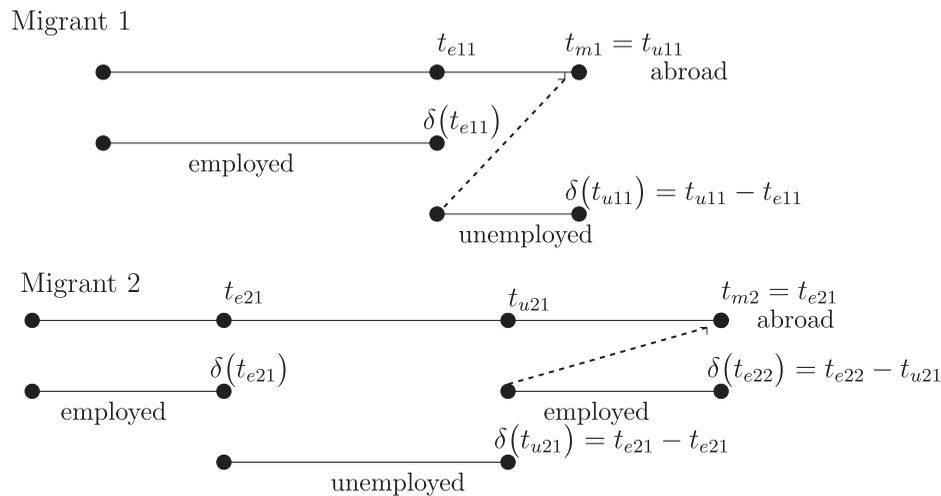
In table 2 we condition on being unemployed at the time of the return migration. In line with the results of table 1, immigrants from the EU15 have a higher incidence of unemployment at the time of their departure (54%), a higher incidence of repeated unemployment (20%), and they are more likely on average to experience longer periods of unemployment (15 months). Immigrants from non-EU DCs have, compared to Europeans, a lower incidence of unemployment (41%) and of repeated unemployment (8%), while their preceding periods of employment were longer on average (20 months).

In table 3 we consider immigrants who, after a period of unemployment, found a job and subsequently left. Hence this group has a volatile labor market experience (employment, followed by unemployment, followed by reemployment), but the last labor market experience is a positive one. Unsurprisingly, the incidence of such labor market histories is low, ranging between 3% and 10%. Although non-European immigrants from DCs exhibit the lowest incidence (3%), the durations of the last two labor market spells look fairly similar across all groups, except for the new EU immigrants, who experience typically shorter durations.

III. The Econometric Model

We seek to determine the effect of labor market dynamics on the return migration intensity of immigrants, so the random outcome variable of interest is the time spent in the Netherlands, denoted by T_m . The observational units are labor immigrants in the host country (the Netherlands). For expositional clarity, we first present a restricted model of the migration period that ignores the labor market processes. These are introduced subsequently, and we address the empirical challenge that arises from the potential correlation between the labor market process and the migration process, which confounds the direct effect of interest.

FIGURE 2.—MIGRATION AND LABOR MARKET DYNAMICS



We follow common practice in duration analysis and express the distribution of the migration duration variate T_m in terms of the associated hazard, say, θ_m . The proportional hazard (PH) model expresses this return hazard as the product between a baseline hazard, $\lambda_m(t)$, which is a function of time alone and common to all individuals, and a covariate function, $\exp(x_m(t)\beta_x^m)$, which accelerates exits: $\theta_m(t|x_m(t)) = \lambda_m(t) \exp(x_m(t)\beta_x^m)$. The covariate vector $x_m(t)$ is allowed to change over time. To accommodate unobserved heterogeneity, the mixed proportional hazard model (MPH) extends the PH model by multiplying it by a time-invariant person-specific error term, say v_m , with distribution G : $\theta_m(t|x_m(t), v_m) = v_m \lambda_m(t) \exp(x_m(t)\beta_x^m)$.

We extend this simple framework in two ways. First, we introduce the MPH models for the unemployment and reemployment processes and describe the timing of events. All three processes are allowed to exhibit correlated effects that can arise from the correlation between the unobserved heterogeneity terms (stemming from, for example, dynamic sorting). We then focus on the direct effect of interest, and consider and discuss the identification argument.

A. Labor Market Processes and the Timing of Events

Generically, let T denote the random time since first entry into the Netherlands that an event takes place. In particular, T_m is the time the immigrant leaves the host country in order to return to the sending country, T_e the time a period of employment ends in the host country, and T_u the time an unemployment period ends. The durations of the periods of employment and unemployment are denoted by $\delta_e(t)$ and $\delta_u(t)$, respectively.

The timing of events and our definitions are illustrated in figure 2. We depict the labor market and migration durations of two migrants. In accordance with our data definitions of section II, migrants are employed at the moment they enter the country. Migrant 1 arrives after migrant 2. The length of migrant 1's (first) period of employment is $\delta(t_{e11}) = t_{e11}$. He remains in the country unemployed until time t_{m1} . His

duration of unemployment is thus $\delta(t_{u11}) = t_{u11} - t_{e11}$. The unemployment is terminated at the moment he leaves the host country at time t_{m1} . Migrant 2 stays longer in the country, $t_{m2} > t_{m1}$, and undergoes a different labor market experience. His first employment has duration $\delta(t_{e21}) = t_{e21}$. After an unemployment period of length $\delta(t_{u21}) = t_{u21} - t_{e21}$, he becomes employed again. This second employment period is terminated when he leaves the host country at time t_{m2} and has duration $\delta(t_{e22}) = t_{e22} - t_{u21}$. The last labor market period for each migrant is always censored. While Migrant 1 experiences an adverse labor market shock (unemployment), migrant 2 experiences a positive shock (reemployment). We seek to determine the effect of such shocks in terms of both their incidences and their durations, on the duration of the migration spell.

Because the migrant is either employed or unemployed, the labor market process is alternating and has three possible transitions: unemployment to employment, employment to unemployment, or return migration. The conditional hazards for periods of unemployment and reemployment follow MPH models:

$$\theta_u(\delta_u(t)|x_u(t), v_u) = v_u \lambda_u(\delta_u(t)) \exp(x_u(t)\beta_x^u), \quad (1)$$

$$\theta_e(\delta_e(t)|x_e(t), v_e) = v_e \lambda_e(\delta_e(t)) \exp(x_e(t)\beta_x^e), \quad (2)$$

with baseline hazards λ_k , unobserved time-invariant characteristics v_k , and observed time-varying characteristics x_k , where $k \in \{u, e\}$ denotes the labor market state. In order to keep track of labor market events, we also define the associated time-varying indicators: the indicator $I_u(t)$ takes a value of 1 if the migrant is unemployed at time t , and $I_r(t)$ indicates that the immigrant is employed again after a period of unemployment. This setting is further generalized by distinguishing between the number of the labor market spell.

The return migration hazard is also of the MPH form. We allow T_m , $\delta_e(t)$, and $\delta_u(t)$ to be correlated through unobservable heterogeneity terms and through a possible direct effect of labor market dynamics on the migration hazard. The latter

is the effect to which we now turn. We consider both the incidence of the unemployment and reemployment event, as well as its duration, and allow the impact to vary systematically with observed characteristics. In our empirical implementation, these three aspects of the impacts (incidence, duration, and heterogeneity) are considered progressively. Thus, the extended MPH model for the return hazard is

$$\begin{aligned} \theta_m(t|t_u, t_e, x_m(t), z(t), v_m) &= v_m \lambda_m(t) \\ &\times \exp\left(x_m(t)\beta_x^m + I_u(t)\left\{\gamma_u + \alpha_u(\delta_u(t)) + z_u(t)\phi_u\right\}\right. \\ &\left.+ I_r(t)\left\{\gamma_e + \alpha_e(\delta_e(t)) + z_e(t)\phi_e\right\}\right). \end{aligned} \quad (3)$$

The covariates z in the (return) migration hazard are a subset of the time-varying characteristics of the migrants x . The duration impacts, α_k , are modeled by piecewise constant functions, so these effects are allowed to exhibit duration dependence. The unobserved error terms (v_e, v_u, v_m) are distributed according to some distribution function G and induce correlated effects, to which we now turn.

B. Endogeneity: Confounding Unobservable Heterogeneity

It is well known that due to dynamic sorting effects, the distribution of v_e among those who become unemployed at t_e will differ from its population distribution. In particular, individuals with high v_e will tend to enter unemployment earlier than individuals with low v_e . If v_e and v_m are dependent, the distribution v_m for unemployed migrants at a given time in the country will differ from the distribution of v_m for migrants still employed. Similarly, if v_m and v_u are not independent, then the distribution of v_m among reemployed migrants will differ from its population distribution. Therefore, one cannot infer the causal effect of unemployment on the return migration from a comparison of the realized durations of those who became unemployed at t_e with the rest of the population, because one would then mix the causal effect of unemployment on the duration with the difference in the distribution of v_m between these migrants. In this case, $I_u(t)$ and $I_r(t)$ will be endogenous, and T_u, T_e , and T_m should be modeled jointly to account for the dependence of the unobserved heterogeneity terms. Therefore, we allow v_u, v_e , and v_m to be correlated.

For parsimony,⁷ we assume that each of the unobserved heterogeneity terms remains the same for recurrent durations of the same type, and we adopt a two-factor loading model with two independent fundamental factors, W_1 and W_2 , both having a discrete distribution on $(-1, 1)$ with

$p_j = Pr(W_j = 1)$. Hence,

$$v_k = \exp(\alpha_{k1}W_1 + \alpha_{k2}W_2), \quad (4)$$

with $k = \{u, e, m\}$. Let $W = (W_1, W_2)'$, $v = (v_e, v_u, v_m)'$, and A be the matrix of factor loadings with rows $A_k = (\alpha_{k1}, \alpha_{k2})$. Note that this two-factor model is very general as it allows for positive and negative correlations among the unobserved heterogeneity terms. The variance-covariance matrix of the unobserved heterogeneity terms is given by $\text{Var}(\ln(v)) = A\text{Var}(W)A'$.⁸ Using a two-point mass specification to model unobserved heterogeneity is common in duration analysis, and extending this to correlated processes by considering a factor-loading specification has also been pursued recently in Crépon, Ferracci, and Fougère (2012) and Osikominu (2013).

C. Identification and the Timing-of-Events Method

The full effect of employment dynamics on remigration hazards is given in our framework by the functions $\gamma_u + \alpha_u(\delta_u(t)) + z_u(t)\phi_u$ and $\gamma_e + \alpha_e(\delta_e(t)) + z_e(t)\phi_e$. An application of the timing-of-events method of Abbring and Van den Berg (2003) implies that these effects can have a causal interpretation. This method requires that the durations of employment, unemployment, and migration are modeled parametrically as MPHs, as we have, and that the so-called no-anticipation assumption holds. Denote by t_e the time an unemployment event would start, and consider first the migration hazard at a time t before the unemployment event. The (untestable) no-anticipation assumption requires that migrants do not anticipate the unemployment event by migrating before the anticipated event would occur. The migration intensity $\theta_m(t|t_e, t_u, x_m(t), v_m)$ is assumed to be affected only for $t > t_e$:

$$\theta_m(t|t_{e1}, \cdot) = \theta_m(t|t_{e2}, \cdot) \quad \text{for all } t < \min\{t_{e1}, t_{e2}\}. \quad (5)$$

We proceed to discuss some potential threats to the validity of the no-anticipation assumption with particular reference to the Dutch case (see the online appendix A for a more detailed discussion of Dutch labor and residency law). The latent risk group consists of currently employed immigrants who expect to lose their jobs at some future time $t_e > t$, and “anticipators” are those among them who migrate at time $t < t_e$. By choosing to migrate, they forgo two payoffs: (a) they terminate their employment voluntarily and thus forgo the expected income flow until time t_e and any redundancy compensation if eligible, and (b) they forgo the option of on-the-job search for a new job before t_e and the option of job search from unemployment after t_e . However, these option values are usually nontrivial. In particular, immigrants who become unemployed are permitted to search for a new job subject to time limits, and table 1 has shown that reemployment probabilities are significant. As for on-the-job search,

⁷ After five years of continuous full-time employment, immigrants in the Netherlands can apply for permanent residency, making them eligible for welfare and employment rights similar to those enjoyed by natives (by contrast, EU15 immigrants gain these almost immediately; see online appendix A for details). Such a status change could imply that the unobserved heterogeneity vector v changes at this date. However, our Kaplan Meier estimates of figure 1 reveal that most immigrants return before month 60, and that at this date, the return probability does not exhibit a jump.

⁸ One additional restriction is needed for identification. We let $\alpha_{m2} = 0$.

external estimates for Dutch data suggest that job-to-job transitions exceed job offer arrival rates for the unemployed (van den Berg & Ridder, 1998, table VII). The scope for anticipation is also limited by the flexibility of the Dutch labor market: unions have little power, and the statutory notice period for a job termination, one month, is short; hence, the potential anticipation period $t_e - t_m$ is likely to be short relative to the migration duration t_m (recall from figure 1 that only a small proportion of immigrants leave within one year and only about 20% have left within two years). In order to gauge the magnitude of the potential bias induced by anticipation, online appendix B considers extensive simulation evidence.

We consider two other potential threats to identification peculiar to the Dutch legal framework: (a) in month 36 of continuous full-time employment any full-time work contract⁹ essentially becomes permanent, and (b) in month 60 of continuous full-time employment, non-EU immigrants gain residence and benefit rights. Some employers might thus have an incentive to terminate a contract at these critical dates. To assess the empirical relevance of these threats, we reconsider the Kaplan Meier estimates of the return probabilities of figure 1. An active systematic threat should manifest itself in terms of jumps. The figure, however, does not suggest the presence of jumps around months 36 and 60.

D. Administrative Removal

Some migrants do not officially inform the authorities that they are about to leave the host country. However, all citizens (immigrants and natives) are required to register with their municipalities (this is a prerequisite for many social services and for tax benefit matters). It is thus clear that any migrant who has no entries in the tax benefit register and does not appear in the register of another municipality must have left the country. Only the exact date of the departure is unknown. Such noncompliers are periodically identified and removed from the registers by the authorities in a step labeled “administrative removal.”

We address this as follows. We assume that the two observable events of administrative removal and zero income at last observed time imply that the migrant has left before the date the administrative removal is recorded, and after the last date of any observed change in the observed characteristics (e.g., labor market status, housing, and marital status). Such limited information is equivalent to interval-censored data. For interval-censored data, the exact end of a duration is unknown, but it is known that the duration ended at some point. If a migrant is administratively removed at duration t_a and the last observed change for this migrant occurred at duration $t_1 < t_a$, the contribution to the likelihood (of the out-migration) of this migrant is the probability of survival until t_1 times the probability that the migrant left the country between t_1 and t_a . The latter is equal to the survival from t_1 until t_a

⁹Unfortunately, we do not have data about the status of an individual’s work contract; however, external estimates (Så, 2008) suggest that the share of temporary contracts among recent male immigrants is about 32%.

given survival. Consequently, administrative removal has no effect on the employed part of the likelihood function or on the likelihood of migrants who are administratively removed with nonzero income until their administrative removal date.

Let a_i indicate whether the emigration of migrant i was due to an administrative removal ($a_i = 1$). For an administratively removed migrant, we introduce two different event dates: t_i^a is the administrative removal date, and $t_i^1 < t_i^a$ is the date of the last recorded change in any of the characteristics of migrant i before t_i^a .

E. Likelihood Function

We have data for $i = 1, \dots, n$ immigrants entering the Netherlands in our observation window. Let K_{ie} and K_{iu} denote the number of the observed employment and unemployment spells of individual i . Note that for some migrants, $K_{iu} = 0$ (e.g., a migrant who remains employed). We consider only the first period of migration. The three indicators Δ_{ik}^u , Δ_{ik}^e , and Δ_i^m signal that k th employment/unemployment, or the migration spell is uncensored. Thus, the likelihood contribution of migrant i conditional on the unobserved heterogeneity $v = (v_e, v_u, v_m)$ is, in the light of the preceding discussions:

$$\begin{aligned}
 L_i(v) = & \prod_{k=1}^{K_{iu}} \left\{ \left[\theta_u(\delta_u(t_{ik}) | \cdot, v_u)^{\Delta_{ik}^u} \right. \right. \\
 & \times \exp\left(-\int_0^{\delta_u(t_{ik})} \theta_u(\tau | \cdot, v_u) d\tau\right) \left. \right]^{(1-a_{ik})} \\
 & \cdot \left[\exp\left(-\int_0^{\delta_u(t_{ik}^1)} \theta_u(\tau | \cdot, v_u) d\tau\right) \right. \\
 & \left. \left. - \exp\left(-\int_0^{\delta_u(t_{ik}^a)} \theta_u(\tau | \cdot, v_u) d\tau\right) \right]^{a_{ik}} I_u(t_{ik}^-) \right. \\
 & \times \prod_{j=1}^{K_{ie}} \left[\theta_e(\delta_e(t_{ij}) | \cdot, v_e)^{\Delta_{ij}^e} \exp\left(-\int_0^{\delta_e(t_{ij})} \theta_e(\tau | \cdot, v_e) d\tau\right) \right]^{I_e(t_{ij}^-)} \\
 & \times \left[\theta_m(t_i | \cdot, v_m)^{\Delta_i^m} \exp\left(-\int_0^{t_i} \theta_m(\tau | \cdot, v_m) d\tau\right) \right]^{(1-a_i)} \\
 & \cdot \left[\exp\left(-\int_0^{t_i^1} \theta_m(\tau | \cdot, v_m) d\tau\right) \right. \\
 & \left. \left. - \exp\left(-\int_0^{t_i^a} \theta_m(\tau | \cdot, v_m) d\tau\right) \right]^{a_i}. \tag{6}
 \end{aligned}$$

This likelihood naturally separates unemployment, employment, and migration periods, and for each allows for censoring and administrative removal. To simplify notation, we have suppressed the dependence on observed characteristics in the hazard rates. $I_u(t_{ik}^-)$ indicates that the migrant is unemployed just before t_{ik} , and similarly for $I_e(t_{ij}^-)$. When $K_{iu} = 0$, the relevant term becomes 1. Note that the last, and only the last, labor market spell is censored. This is either because the migrant is still in the country at the end of the observation period or has left.

Integrating out the unobserved heterogeneity, we obtain the likelihood function

$$L = \prod_{i=1}^n \int \int \int L_i(v) dG(v_e, v_u, v_m), \tag{7}$$

where $G(v_e, v_u, v_m)$ is the joint distribution of the unobserved heterogeneity terms implied by the discussion of v_k given by equation (4).

IV. Results

The full model, given by the MPH hazards (1) to (3) and the heterogeneity model (4), permits the effects of the labor market processes to exhibit duration dependence and vary systematically with observed characteristics, while controlling for correlated effects that arise from correlated unobserved heterogeneity. In view of this complexity, and given the nested structure of the model, we first consider simpler variants before turning to the full model.

A. The Restricted Model

In order to gain a first descriptive impression of the effect of the incidence of unemployment on migrants' return, we compare the Kaplan Meier estimates for the group of immigrants who always work and those who have experienced unemployment once (at the date of their return, the latter might be either still unemployed or are reemployed). For short durations, this comparison is, of course, distorted, as immigrants who experience changes in their labor market status must have stayed for a sufficiently long period to allow such events to take place.

Figure 3 depicts the Kaplan Meier estimates for all immigrants (the plots for each immigrant group look similar). It is evident that for durations above 20 months, unemployment seems to have a significant effect on return probabilities. The size of the effect seems to increase in the duration of stay. We proceed from this descriptive analysis to the analysis based on our empirical model.

We focus first on the incidence of labor market events by setting the varying duration and heterogeneity impacts to 0, $\alpha_k = \phi_k = 0$, and refer to this as the timing-of-events model 1. It nests the PH model that ignores unobservable heterogeneity altogether, $\theta_m^{PH}(t|t_u, t_e, x_m(t)) = \exp(x_m(t)\beta_x^m + I_u(t)\gamma_u + I_r(t)\gamma_e)$, whereas the MPH model, $\theta_m^{MPH}(t|t_u, t_e, x_m(t), v_m) = v_m \theta_m^{PH}(t|t_u, t_e, x_m(t))$ ignores the correlation between λ_m and (λ_u, λ_e) . The coefficients of interest are γ_u and γ_e .

Table 4 reports the estimates of the effects given by γ_u and γ_e . The simple PH model already demonstrates the importance of labor market events on return migration. In all cases (except one), the incidence of an unemployment spell significantly increases return probabilities, and the event of finding employment increases migration durations. Extending this model to incorporate (uncorrelated) unobserved heterogeneity has only a small effect on the estimated causal effect.

FIGURE 3.—KAPLAN MEIER ESTIMATES OF RETURN PROBABILITIES BY LABOR MARKET EXPERIENCE

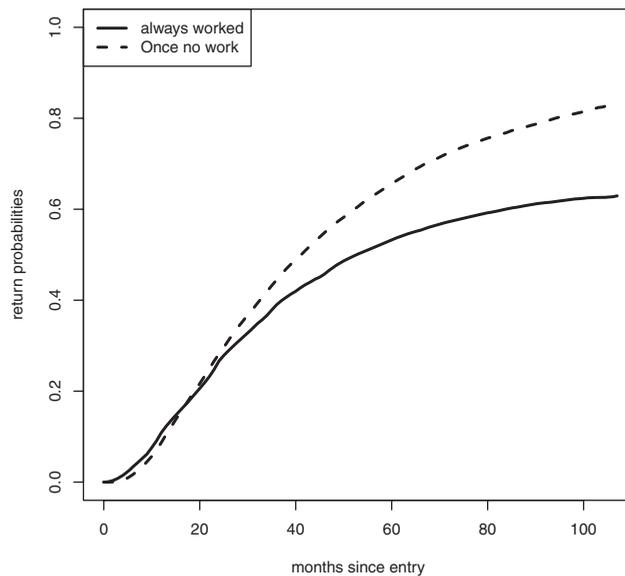


TABLE 4.—INCIDENCE IMPACT OF LABOR MARKET EVENTS ON RETURN-MIGRATION HAZARDS

	EU 15	New EU	Non-EU	
			DCs	LDCs
<i>Unemployment</i>				
PH model	1.045*** (0.095)	0.938*** (0.192)	0.103 (0.219)	1.400*** (0.300)
MPH model	1.144*** (0.099)	0.967*** (0.193)	0.267 (0.239)	1.501*** (0.307)
Timing-of-events model 1	0.778*** (0.105)	0.684*** (0.197)	0.223 (0.237)	1.207*** (0.314)
<i>Reemployment</i>				
PH model	-0.179*** (0.027)	0.203*** (0.062)	-0.370*** (0.078)	-0.383*** (0.061)
MPH model	-0.163*** (0.029)	0.217*** (0.063)	-0.332*** (0.088)	-0.367*** (0.064)
Timing-of-events model 1	-0.091** (0.038)	0.277*** (0.069)	-0.356*** (0.125)	-0.093 (0.083)

p < 0.05; *p < 0.01.

However, taking into account the correlated unobserved heterogeneity in the timing-of-events model substantially affects the estimates. The estimated effect is typically smaller in magnitude than for both PH and MPH models. For instance, as regards the incidence of unemployment for EU15 immigrants, the MPH estimate is 1.144, whereas our model estimate is .778, the two differing by a factor of 1.47. We conclude that ignoring the endogeneity issue would result in substantial selectivity biases.

B. The Full Model

Building on the insights of the timing-of-events model 1, we first allow the durations of the labor market events to have an impact on immigration durations more flexibly. This is implemented nonparametrically by modeling the sequence $\{\alpha_k\}$ as piece-wise constant functions. Thus, in model 2, we have $\gamma_u = \gamma_e = 0$ and $\phi_u = \phi_e = 0$. In model 3, we further

TABLE 5.—ESTIMATED EFFECT OF BECOMING UNEMPLOYED ON RETURN MIGRATION HAZARDS

	Model 1				Model 2				Model 3			
			Non-EU				Non-EU				Non-EU	
	EU 15	New EU	DCs	LDCs	EU 15	New EU	DCs	LDCs	EU 15	New EU	DCs	LDCs
Constant effect [γ_u]	0.778*** (0.105)	0.684*** (0.197)	0.223 (0.237)	1.207*** (0.314)								
<i>Duration dependence</i> [α_u]												
0–3 months					0.643*** (0.106)	0.585*** (0.202)	0.379 (0.222)	1.261*** (0.308)	0.865*** (0.117)	1.814*** (0.272)	0.108 (0.241)	1.058*** (0.326)
3–6 months					0.658*** (0.107)	0.910*** (0.206)	0.282 (0.224)	1.406*** (0.311)	0.983*** (0.119)	2.054*** (0.275)	0.047 (0.243)	1.183*** (0.328)
6–12 months					0.517*** (0.107)	0.698*** (0.211)	-0.214 (0.226)	1.358*** (0.314)	0.939*** (0.119)	1.752*** (0.277)	-0.423 (0.244)	1.116*** (0.328)
More than 1 year					0.312*** (0.117)	0.748*** (0.226)	-0.509** (0.224)	1.575*** (0.328)	1.009** (0.123)	1.562*** (0.278)	-0.671*** (0.243)	1.315*** (0.336)
<i>Labor market history</i> [ϕ_u]												
Repeated unemployment									-0.198*** (0.057)	-0.523** (0.205)	0.067 (0.232)	-0.459*** (0.145)
Order of unemployment spell									-0.186*** (0.032)	-0.046 (0.126)	-0.216 (0.182)	-0.156 (0.083)
<i>Duration of previous employment spell</i>												
Less than 3 months									-0.302*** (0.043)	-0.235** (0.117)	0.065 (0.122)	-0.016 (0.103)
3–6 months									-0.338*** (0.043)	-0.081 (0.103)	0.116 (0.087)	-0.119 (0.096)
More than 1 year									-0.318*** (0.036)	-0.189** (0.092)	0.032 (0.064)	-0.153** (0.072)

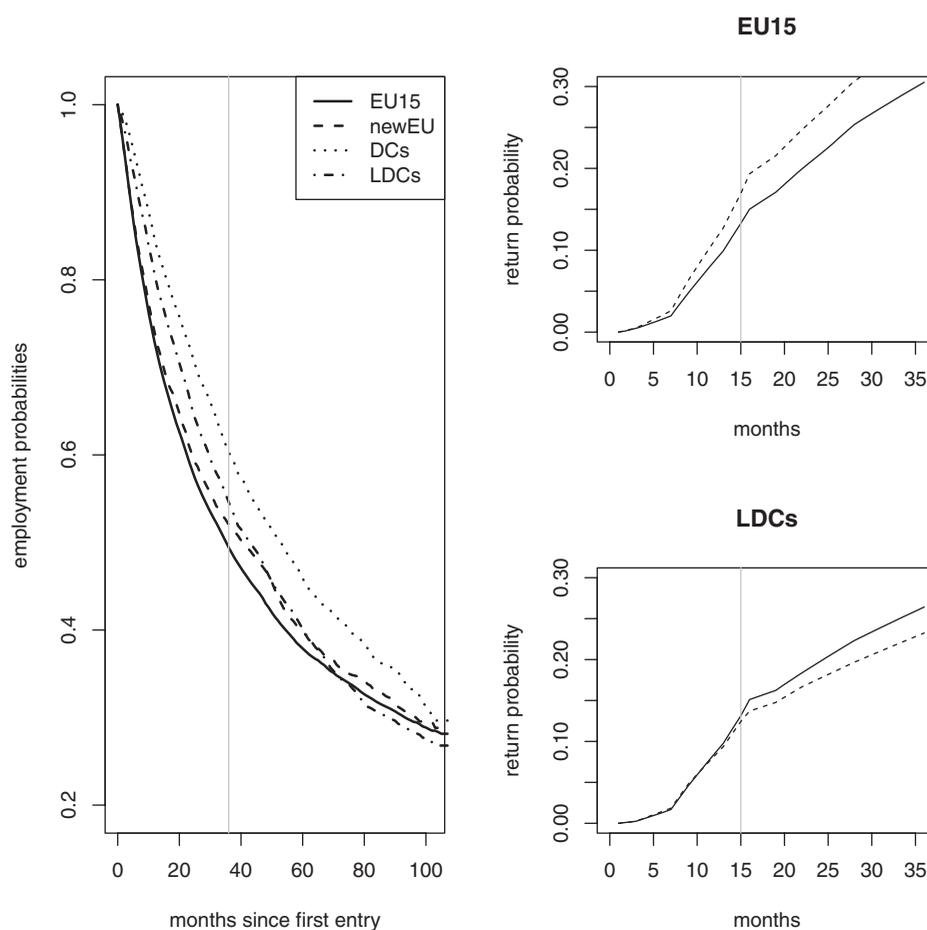
The model equations are given by (1) to (3), the likelihood by (7). SE in parentheses. **: $p < 0.05$ and ***: $p < 0.01$. Model 3 covariates (z) also include demographics (sex, married, number of children, age group dummies, and cohort effects measured by the year of entry). Reference category for employment durations: 6–12 months. "Order of unemployment spell" refers to the second, third, etc. period of unemployment.

TABLE 6.—ESTIMATED EFFECT OF REEMPLOYMENT ON RETURN MIGRATION HAZARDS

	Model 1				Model 2				Model 3			
			Non-EU				Non-EU				Non-EU	
	EU 15	New EU	DCs	LDCs	EU 15	New EU	DCs	LDCs	EU 15	New EU	DCs	LDCs
Constant effect [γ_e]	-0.091** (0.038)	0.277** (0.069)	-0.356** (0.125)	-0.093 (0.083)								
<i>Duration dependence</i> [α_e]												
0–3 months					-0.222** (0.089)	0.298*** (0.113)	-0.233 (0.182)	-0.327** (0.147)	-0.096 (0.082)	0.339 (0.205)	-0.164 (0.253)	-0.302 (0.186)
3–6 months					0.017 (0.088)	0.454*** (0.122)	-0.190 (0.192)	0.000 (0.142)	0.127 (0.083)	0.532** (0.209)	0.132 (0.260)	0.085 (0.185)
6–12 months					-0.042 (0.080)	0.203 (0.119)	-0.097 (0.145)	-0.102 (0.123)	0.083 (0.077)	0.319 (0.211)	0.054 (0.229)	0.038 (0.172)
More than 1 year					-0.315*** (0.058)	0.194 (0.108)	-0.641*** (0.119)	-0.390*** (0.095)	-0.242*** (0.072)	0.357 (0.210)	-0.626*** (0.214)	0.140 (0.158)
<i>Labor market history</i> [ϕ_e]												
Repeated reemployment									0.068 (0.091)	0.327 (0.214)	-0.073 (0.586)	0.261 (0.202)
Order of reemployment spell (>1)									-0.194*** (0.051)	-0.055 (0.125)	-0.318 (0.423)	-0.097 (0.107)
On benefit									-0.431*** (0.085)	-0.825*** (0.314)	0.048 (0.284)	-0.570*** (0.210)
<i>Duration of previous unemployment spell</i>												
Less than 1 month									-0.406*** (0.080)	-0.332** (0.165)	-0.128 (0.285)	-0.281 (0.174)
1–2 months									-0.284*** (0.079)	-0.138 (0.173)	0.298 (0.248)	-0.580*** (0.195)
2–3 months									-0.185** (0.080)	0.100 (0.168)	-0.037 (0.252)	-0.164 (0.172)
6–12 months									0.030 (0.080)	-0.179 (0.188)	0.294 (0.226)	-0.279 (0.184)
More than 1 year									0.079 (0.101)	-0.162 (0.238)	-0.444 (0.305)	-0.218 (0.240)

See table 5 notes.

FIGURE 4.—ASSESSING THE IMPACT OF TRUNCATING IMMIGRATION DURATIONS AT MONTH 36



Left: Kaplan Meier estimates of employment probabilities. Right: Impacts on return probabilities for unemployment periods of 12 months that start in month 3 (broken lines are based on the truncated data at month 36; the solid lines are based on the unrestricted data).

allow the effect to be heterogeneous across migrants in terms of z_u and z_e , which measure demographics and previous labor market history. For expositional clarity, we present the effect of the periods of unemployment and reemployment spells in tables 5 and 6 separately, although it is clear from equation (7) that these are estimated simultaneously. For ease of reference, model 1 reports the results of the last row of table 4. For brevity, we do not discuss the coefficients of the covariates x_k which are of only secondary importance.¹⁰

C. The Effects of Becoming Unemployed

The estimated effects of unemployment spells on return migration hazards are reported in table 5. Across all three specifications and all immigrant groups, it is evident that unemployment spells shorten migration durations.¹¹

¹⁰ The estimates are available from the authors. The covariates include extensive measures of demographics, a nonparametric function of income, housing descriptors, sector dummies, cohort effects measured by the year of entry, and controls for macro effects.

¹¹ In the discussion paper version of this paper (Bijwaard et al., 2012), we quantify the impacts of the unemployment process in various evaluation exercises in terms of the return probability $1 - Pr\{T_m > t\}$ where

We recall that the average effect γ_u for all of the groups of immigrants estimated in model except for immigrants from non-EU DCs, the significant point estimates ranging from 0.68 to 1.2. Models 2 and 3 reveal that the effect exhibits duration dependence. In model 2, for EU migrants, the impact peaks for durations of 3 to 6 months. For non-EU migrants, the picture is more heterogeneous, as duration dependence increases for immigrants from LDCs, whereas the coefficients remain insignificant for the others except for long durations. In model 3, we further permit the effect to vary across characteristics (demographics and labor market history). This increases the magnitude of the duration effects for EU immigrants, and follows since the duration of the preceding employment spell lengthens the migration spell. By

$Pr\{T_m > t\} = E_{v_m} \{ \exp(-\int_0^t \theta_m(s) ds) \}$. For instance, consider a reference individual ($\beta_x^m = 0, \phi_u = 0, \phi_e = 0$) from the EU15 who starts a period of unemployment three months after entry into the Netherlands in 2003. If he is unemployed for twelve rather than three months, the maximal increase in the return probability is 6 percentage points. If he is from an LDC, the maximal difference is 14 percentage points. Holding the length of the unemployment period constant at six months, but starting it progressively later, we found a maximal increase in return probabilities of 6 percentage points.

TABLE 7.—SENSITIVITY ANALYSIS: MIGRATION DURATIONS LESS THAN THREE YEARS

	Model 1				Model 2				Model 3			
	EU 15	New EU	Non-EU		EU 15	New EU	Non-EU		EU 15	New EU	Non-EU	
			DCs	LDCs			DCs	LDCs			DCs	LDCs
Estimated Effect of Unemployment on Return Migration Hazards												
Constant effect [γ_u]	0.574*** (0.129)	0.665*** (0.213)	0.324 (0.280)	1.234*** (0.378)								
<i>Duration dependence</i> [α_u]												
0–3 months					0.700*** (0.120)	0.570*** (0.218)	0.078 (0.257)	1.180*** (0.373)	0.756*** (0.135)	1.837*** (0.326)	-0.160 (0.282)	1.015*** (0.394)
3–6 months					0.818*** (0.124)	0.893*** (0.223)	0.007 (0.259)	1.223*** (0.377)	0.861*** (0.138)	2.101*** (0.329)	-0.215 (0.284)	1.014*** (0.396)
6–12 months					0.721*** (0.127)	0.645*** (0.227)	-0.561* (0.264)	1.049*** (0.381)	0.766*** (0.140)	1.805*** (0.335)	-0.758*** (0.290)	0.789** (0.398)
More than 1 year					0.826*** (0.139)	0.666*** (0.246)	-0.781*** (0.268)	1.007*** (0.395)	0.946*** (0.149)	1.713** (0.354)	-1.010*** (0.300)	0.632 (0.406)
Estimated Effect of Reemployment on Return Migration Hazards												
Constant effect [γ_e]	-0.008 (0.051)	0.341*** (0.080)	0.310 (0.176)	-0.071 (0.100)								
<i>Duration dependence</i> [α_e]												
0–3 months					-0.163** (0.076)	0.329*** (0.122)	0.188 (0.227)	-0.101 (0.166)	-0.025 (0.103)	0.576*** (0.223)	-0.079 (0.321)	-0.027 (0.239)
3–6 months					0.091 (0.074)	0.560*** (0.131)	0.309 (0.246)	0.116 (0.170)	0.244** (0.103)	0.814*** (0.227)	-0.003 (0.330)	0.259 (0.242)
6–12 months					0.039 (0.065)	0.219 (0.140)	0.460** (0.215)	-0.149 (0.166)	0.209** (0.098)	-0.474** (0.235)	0.082 (0.308)	0.048 (0.240)
More than 1 year					-0.339*** (0.072)	0.211 (0.182)	0.288 (0.268)	-0.270 (0.173)	-0.151 (0.106)	0.488 (0.262)	-0.105*** (0.330)	0.019 (0.252)

See the notes to table 5.

contrast, the effect of the previous labor market history is found to be insignificant for non-EU DC immigrants.

D. *Effects of Becoming Reemployed*

Finding employment after a period of unemployment is a positive labor market event that is likely to have an impact on migration duration. Table 6 reports the results. For all except non-EU LDC immigrants, the effect of having found employment after being unemployed delays the migrant’s return. The effect is particularly strong for immigrants from developed countries outside the EU. Previous unemployment durations exhibit an effect only if these were no longer than three months, indicating that such periods of unemployment were anomalies that the individual quickly overcame. The one immigrant group that deviates from this pattern of extended migration durations is made up of immigrants from the new EU—mainly Polish immigrants. The estimated impact of reemployment for this group, however, is consistent with target savings: having regained employment, it is plausible that such immigrants are back on track to reach their savings target and return once they have saved enough.

E. *Sensitivity Checks*

Our modeling approach assumes that unobserved heterogeneity v is time invariant. One concern is that this might not be valid for individuals who experience multiple labor market transitions. However, we cannot condition on the experience of only one period of unemployment because of the ensuing

selection bias. We address this concern by truncating the data at 36 months since entry, a period sufficiently long to allow labor market events to take place while at the same time being sufficiently short to make multiple transitions less likely.

The left panel of figure 4 depicts the estimates of the Kaplan Meier estimates of unemployment probabilities and the imposed cut-off date. The plot illustrates, as did our earlier descriptive tables, the substantial unemployment risk that all immigrant groups experience and the importance of stratification by immigrant group. Table 7 reports, for these restricted data, the estimates of models 1 to 3, concentrating for reasons of space on the estimates of the impact coefficients γ_k and α_k .

The results for the restricted durations are similar to those for the unrestricted sample of tables 5 and 6. The coefficients have changed only slightly without affecting the qualitative conclusions. The right panel of figure 4 illustrates the effect of the change in the coefficients brought about by the truncation at month 36. These are assessed in terms of the effect on return probabilities for the representative immigrant, assuming an unemployment period of twelve months that starts in month 3 (see also note 11). For EU15 immigrants, there is a small effect on return probabilities for the higher durations, whereas for LDC immigrants, there is no discernable difference.

Another concern could be our selection rule that the immigrant be employed within three months of entry into the Netherlands, imposed to limit the potential of misclassification of the labor motive. We have reestimated the models for varying cut-offs. For brevity, table 8 reports the estimates of the impacts of the labor market process on migration hazards

TABLE 8.—EFFECT OF CHANGING THE INITIAL EMPLOYMENT REQUIREMENT IN MODEL 1

	EU 15	New EU	Non-EU	
			DCs	LDCs
<i>Unemployment</i>				
2 months	0.759***	0.607***	0.171	1.232***
3 months	0.778***	0.684***	0.223	1.207***
4 months	0.767***	0.714***	0.213	1.241***
5 months	0.739***	0.703***	0.280	1.248***
<i>Reemployment</i>				
2 months	-0.078**	0.273***	-0.372***	-0.080
3 months	-0.091**	0.277***	-0.356***	-0.093
4 months	-0.099***	0.282***	-0.463***	-0.120
5 months	-0.105***	0.296***	-0.443***	-0.095

** $p < 0.05$; *** $p < 0.01$.

of model 1 (the same conclusions obtain for models 2 and 3). We observe only a small change in the point estimates, and all qualitative conclusions remain unchanged.

Online appendix B contains several other simulation-based robustness checks: we verify the robustness of our estimation procedure, conduct a simulation-based placebo test assuming no labor market impact, and examine the sensitivity of the results to departures from the no-anticipation assumption.

V. Conclusion

The majority of recent labor immigration to the Netherlands is temporary rather than permanent. Across all immigrant groups, a substantial proportion leave the host country eventually, and many do so within 24 months. We have considered in this paper the individual labor market drivers of immigration durations.

Despite this extent of temporary immigration, the interdependence of labor market events and immigration durations has received little attention in the empirical literature. We have addressed this gap using a unique Dutch administrative panel of the entire population of recent labor immigrants, in which we observe entry, exit, reason for migration, and complete labor market histories. The large size of the data permitted us to stratify the analysis by distinct immigrant groups. The timing-of-events method enabled us to estimate the effects of employment and unemployment histories on migration durations, while we controlled for unobservable heterogeneity. Simpler models that ignore such error

correlations across labor market and migration processes were shown to exhibit substantial selection biases.

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