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Gender differences in labour market integration trajectories of recently arrived migrants in the Netherlands

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ABSTRACT

This study investigates gender differences in recently arrived migrants' labour market activity and occupational status both shortly after arrival and with increasing length of stay. We examine the role of education, household composition and traditional gender role values by estimating multi-group multilevel models based on three waves of the New Immigrants to the Netherlands Survey. In line with findings regarding gender gaps in labour market behaviour, recent female migrants are less active on labour market than their male counterparts, and we observe a clear motherhood penalty and fatherhood premium on the number of hours worked. Men and women show only marginal differences in their occupational statuses. Changes over time do not differ between men and women, indicating persistent gender inequality in labour market attainment. Moreover, interesting differences between the nationalities were found. Polish migrants show the highest activity levels and lowest occupational status, also when compared to Bulgarians. Spanish migrants hold the highest occupational statuses. Recent Turkish migrants seem to be better integrated and show fewer gender differences than the more established Turkish minority in the Netherlands.

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Labour market integration; recently arrived migrants; female labour market participation; longitudinal design; multilevel multi-group model

Introduction

Findings according to which migrants have unequal labour market integration patterns depending on their gender (e.g. Adsera and Chiswick 2007; Alesina and Giuliano 2010) are alarming, since successful integration into the labour market indicates and reinforces integration in other areas as well (Foroutan 2008). In addition to economic benefits, being employed means, for example, that migrants have more opportunities for contact with the native population (Martinovic, Van Tubergen, and Maas 2009) and become more exposed to the local language (Chiswick and Miller 2001). The time shortly after migration is considered as a critical stage for the future integration process (Chiswick and Miller 1994; Busk et al. 2016), but quite surprisingly, the labour market trajectories of migrants who have recently arrived have been scarcely studied.[1] It is therefore not clear whether gender differences in labour market integration are already visible upon arrival, or emerge during the integration process. Longitudinal data that allow to study accurately

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how the labour market trajectories change over time, and especially over the first years after migration, are needed to answer this question, and this is what this study can offer.

Human capital theory is often used to understand migrants' labour market integration (Chiswick 1978; Becker 1985). Human capital consists of, for example, skills and education, with a greater stock predicting better labour market positions. However, even among the highly educated, marriage and children have disproportionally more negative effects on women's employment (Alesina and Giuliano 2010). One explanation for this is that people believe in traditional gender roles according to which women have the responsibility of taking care of children and home, whereas men have the economic responsibility for the family (Alesina and Giuliano 2010; Stam, Verbakel, and de Graaf 2014). The extent to which people believe in these gender roles can further explain within-gender differences in labour market participation (McRae 2003). Traditional gender role values are considered to be more pronounced outside western countries, and milder among every generation that has migrated across this border of egalitarianism (Röder 2014).

The current study contributes to knowledge about recently arrived migrants' labour market integration. We study how initial positions differ between men and women and how they change during the first years after migration. As it is known that the labour market position and gender gaps therein also depends on migrants' country of origin (Fleischmann and Höhne 2013), we study four distinct origin groups to enhance the generalisability and external validity of our conclusions.

Recent Immigrants in the Netherlands

We study differences among recent migrants from Bulgaria, Poland, Spain and Turkey in the Netherlands, which is one of the most important migration-destination countries in Europe (Crul and Heering 2008). A great share of the migrant population is originally either from old colonies (Antilles and Surinam), or from Turkey and Morocco, where so-called guest workers were recruited in the 1960s. Even though much research has already been conducted regarding the more established Turkish minority in the Netherlands, it is important to study whether the characteristics that are often found among them, i.e. comparatively weak labour market positions (Zorlu 2013) and low participation rates among women (Khoudja and Fleischmann 2015), also apply to more recent Turkish migrants. Bulgarian, Polish and Spanish migrants do not have as long roots in the country as the Turkish (Bertoli, Brücker, and Moraga 2013) but are among the largest groups of new migrants coming to the Netherlands (OECD 2016). Unlike Turkish migrants, they are migrating from another EU country, which means that they have freedom of movement (Galgóczi and Leschke 2016). The labour market participation trajectories of these groups are less well known and thus investigated in the current study.

Labour market integration

Before going more into detail regarding theories of labour market integration, we shortly note that the concept of labour market integration is operationalised – and hence measured – differently by various scholars. We analyse both the *working hours* and *occupational status* to study labour market integration.

Labour market integration and working hours

In line with human capital theory, higher educated persons are usually more active on the labour market, for example, because they have better opportunities for being employed and they receive higher returns for their activity (e.g. Adsera and Chiswick 2007). We expect that also among recently arrived migrants the more highly educated work more hours (H1a). Furthermore, we expect that higher educated migrants fluctuate less in their labour market activity over time (H1b). We argue that this is both because they have less room and less need to increase their working hours if they already work relatively many hours shortly after arrival. What is important, especially among migrants, men are often more active in the labour market than women, even if they have a similar level of education (e.g. Raijman and Semyonov 1997; Fleischmann and Höhne 2013). We expect to find similar patterns among recently arrived migrants, and hypothesise that men work more hours than women (H2). To understand the underlying reasons for this gender difference, we turn to family composition and traditional gender role values.

Most literature investigating the effect of a partner focuses either on the relative earnings or the social capital of the partner (e.g. Verbakel 2008). However, as we focus on recently arrived migrants, it is likely that their partners are also recently arrived,[2] and thus would have relatively little variation in their host country-specific social capital. Therefore, we focus on the mere presence of a partner. Having a partner is generally found to have a negative effect on women's labour market activity, whereas the opposite is found for men (Raijman and Semyonov 1997; Bevelander and Groeneveld 2012). However, according to the family investment hypothesis (Baker and Benjamin 1997), having a partner might affect the gender differences in labour market integration in *different ways* at *different stages* of the recently arrived migrants' integration trajectories. This hypothesis suggests that when couples migrate together, women are more active on the labour market upon arrival, often in dead-end jobs, while their male partners initially focus on gaining host country-specific human capital. This would lead to the maximisation of household utility in the long run, as it allows the male partner to enter the labour market with improved qualifications, allowing the women to decrease their activity. In sum, the labour market trajectories of male and female migrants would be markedly different if they arrive with a partner, but less so if they migrate individually.

In most cultures, women are seen as natural caretakers for small children and men as providers for the family (Stam, Verbakel, and de Graaf 2014). Women who have children usually show lower levels of labour market activity whereas the opposite is found for men (Kaufman and Uhlenberg 2000). Some studies have also found that having children has no effect on men's working hours (McGill 2014). The negative effect of motherhood is especially likely to hold if families do not have a large social network to help with childcare (Adsera and Chiswick 2007; Banerjee and Phan 2015), which is usually the case among recently arrived migrants who have had little time to establish social networks in their destination country.

The effect of having a partner and children might vary depending on how strongly people believe in traditional gender roles. Women who hold more traditional gender role values have been found to be less active on the labour market (Stam, Verbakel, and de Graaf 2014), and especially so when they have children (Hakim 2000; Khoudja and Fleischmann 2015). The situation is the opposite for men, as traditional gender role values usually further increase fathers' working hours (Kaufman and Uhlenberg 2000).

The negative consequences of having partner and children are often smaller for women who are highly educated, since the opportunity costs of staying home are higher for them (Bevelander and Groeneveld 2012; England et al. 2016) and they are likely to have more egalitarian gender role values (Judge and Livingston 2008). Highly educated men usually increase their working hours less as consequence of having a partner or children than those who have a lower education, since they have less room for increase (Kaufman and Uhlenberg 2000), and more egalitarian gender role values (Judge and Livingston 2008).

In sum, for recently arrived migrants, we hypothesise that having a partner has a negative effect on women's working hours but a positive effect for men's (H3a), especially when they hold traditional gender role values (H3b) but less so if they are highly educated (H3c). We expect that women who have children work fewer hours (H4a), and especially so if they have strong traditional gender role values (H4b), but less so if they are highly educated (H4c). We test whether traditional gender role values have a direct negative effect on women's working hours (H5) and explore the parallel effect for men. We expect that men with children work more hours (H6a) and especially so if they have more traditional gender role values (H6b) but less so if they are highly educated (H6c).

Labour market integration and occupational status

Occupational segregation based on gender is a persistent global phenomenon (Murphy and Cross 2017), and a lot of research has studied gendered preferences (e.g. Hakim 2000) and wage differences (Hegewisch and Hartmann 2014). The level of education is one of the most important predictors for the level of occupational status (Foroutan 2008), but it also has an impact on how fast one can advance in the status hierarchy. However, this effect is not always equal for men and women: for example, Amuedo-Dorantes and De la Rica (2007) found that among recent migrants in Spain, the highly educated experienced overall steeper occupational mobility, but the returns to education were higher for men than women. Similarly, not only do migrants overall have difficulties in obtaining higher occupational status (e.g. Heath and Cheung 2007), but especially migrant women often hold the lowest occupation positions (Bevelander and Groeneveld 2012). We hypothesise that also among recently arrived migrants the higher educated have higher occupational status (H7a) but that women have lower status than men (H7b). We expect that higher educated migrants improve their occupational status faster over time (H8a) but that the positive effect of education is smaller for women than for men (H8b).

The relationship between having a partner and occupational mobility is rather ambivalent, and most studies mainly focus on partner resources (e.g. Brekke 2013). However, as discussed above, we are only interested in the presence of a partner. One study that investigated the mere presence of partner found no effect on women's occupational mobility, but a small boost for men (Verbakel and de Graaf 2008). Therefore, we hypothesise that among recently arrived migrants, having a partner has a positive effect on occupational status for men but not for women (H9).

The motherhood penalty is found to be relevant also for occupational status (Abendroth, Huffman, and Treas 2014). Staying at home with children can have both short and long-term costs for women's occupational status. They might not seek for more demanding occupations, or employers might be cautious in hiring mothers of small children in more demanding positions, and employers can be reluctant to hire candidates who have stayed home with children for a longer period (Abendroth, Huffman, and Treas 2014). Also here, the negative effect of motherhood is suggested to be lower for women with higher education (Budig and Hodges 2010), since they have greater opportunities, for example, to pay for their children's day-care (Hook and Pettit 2016). However, there are also findings according to which the educational level of the mother does not significantly affect the size of the motherhood penalty (e.g. Berghammer 2014). To our knowledge, the role of traditional gender role values has not been discussed in relation to occupational status in the literature, but we argue that similar effects could be expected as for working hours, since more 'traditional' mothers would, for example, stay home with their children longer. We hypothesise that for recently arrived migrants, having children has a negative impact on women's occupational status (H10a) and that the negative effect of motherhood is smaller for mothers with higher education (H10b) and larger if they believe more strongly in traditional gender role values (H10c).

For men, the relationship between having children and occupational status is mainly studied trough income. Fatherhood has often a positive effect on wages, especially if one is already in relatively high position or has high education (Budig and Hodges 2010). As income is an important indicator of the level of occupational status, we expect that having children has indeed a positive effect on men's occupational status (H11a), which should be even larger among higher educated men (H11b) (Figure 1).

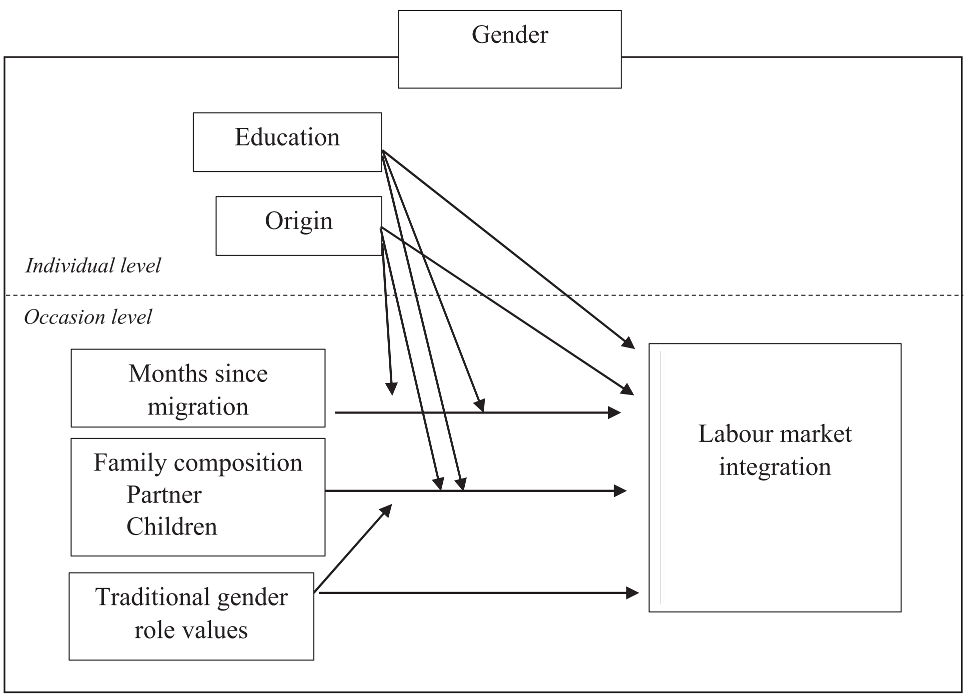

Figure 1. Theoretical model.

Note. All dependent variables and independent variables at the lowest level are treated as time-dependent variables and measured at three different time points. Gender is used as a grouping variable.

Data and methods

Data and participants

This study uses three waves of panel data from the New Immigrant Survey Netherlands (NIS-2NL).[3] The survey resulted in a large-scale longitudinal data-set about the integration processes of recently arrived migrants from Bulgaria, Poland, Spain and Turkey in the Netherlands. Respondents received the survey in their native language and could choose to complete it either online or on paper. The first wave was collected in two rounds (fall 2013 and spring 2014), and all respondents were approached within a year after registering with a Dutch municipality. A total of 4808 respondents filled in the questionnaire at the first wave. For the second and third wave, all respondents who completed the first and if applicable, the second round, who had agreed to participate in the following round and who were still living in the Netherlands[4] were contacted. The second wave was collected in early and mid-2015, and the third wave in late 2016. The average time between waves was 15 months.

Some of the migrants in the survey had already been living in the Netherlands for several years, even though they only registered recently. To focus on recently arrived migrants, respondents who indicated at the first wave that they arrived more than 5 years ago[5] were excluded ($N = 399$). Respondents who provided inconsistent answers for their year of birth and gender across waves were also excluded ($N = 155$). Respondents who did not report their gender and for whom time since migration was missing were excluded from the analysis. The sample used for the analysis consists therefore of 3797 respondents for the first wave, 1752 for the second wave and 1029 for the third wave. Age at the first wave varied from 14 to 69 years, with a mean of 30.24 (SD = 8.17). At the first wave, 55.7% of the respondents were women, and 18.0% of the respondents were from Bulgaria, 34.7% from Poland, 30.3% from Spain and 16.9% from Turkey.[6]

Measures

Labour market integration was assessed with two measures to capture both quantitative and qualitative aspects of labour market integration. First, the self-reported number of *hours worked per week* was utilised. In order to avoid the results being biased by extreme outliers, the measure was truncated at 60 h per week.[7] To differentiate between the reasons why some respondents worked 0 h per week, we controlled for *labour market activity*.[8] This was based on an item asking the respondent's main activity, with answer options being working, unemployed, in education, retired, long-term sick or disabled, looking after home or children, on maternity or paternity leave. From these, we constructed two dummies: in education (also taking into account answers to questions about full-time education) and inactive (including respondents who were unemployed), with working as the reference category. Respondents who said they were working, but did not report working hours *or* occupational status ($N = 2$) were recoded as inactive.

The second outcome variable for labour market integration was *occupational status*. Respondents stated the name or title of their job in response to an open question. These answers were categorised using the International Standard Classification of Occupation (ISCO2008), and for the purposes of the present study, further recoded using the International Socio-Economic Index (ISEI) for occupational status (Ganzeboom and

Treiman 2012). This scale reflects the average level of education and earnings in each occupation, and unlike the ISCO scale, the ISEI scale can be treated as a continuous measure. The range of the ISEI scale varies from 0 to 100, where 100 is highest occupational status. In the current sample, the ISEI score varied between 11.56 and 88.31. Only respondents who worked were considered to have valid answers for occupational status. Of those, status was missing for 41.2% at the first wave, 57.9% at the second wave and 20.3% at the third wave.

Family composition was studied with two different measures, both of which were constructed from a set of questions about respondents' household composition. We measured whether a respondent lived together with a *partner,* and with *children*.[9] Household information was missing for 554 respondents at the first wave, 229 at the second wave, and for nobody at the third wave. Respondents with missing answers on the household composition questions were recoded to have *no* spouse or children. A dummy variable for respondents with missing information was included in the analysis to test the impact of this imputation.

Traditional gender role values were assessed with the level of agreement with three statements: 'It is more important for men to earn their own income than it is for women', 'Decisions about large purchases are best taken by men' and 'Women should stop working after they have children'. The possible answers varied from 1 = strongly agree to 5 = strongly disagree. An exploratory factor analysis extracted a one-factor solution explaining 68.6% of the variance in the observed covariation matrix, and all three items loaded highly (>.79) on the factor. Cronbach's alpha for the one-factor solution was acceptable ($\alpha = .766$ for the pooled sample, for different origin country groups α varied between .685 for Spanish and .772 for Turkish), and therefore we computed a mean score as a measure of *traditional gender role values.* The original scores were reversed so that a higher score indicates more traditional values.

Country-specific *education levels* were recoded following the International Standard Classification of Education ISCED 2011 (OECD/Eurostat/UNESCO Institute for Statistics 2015) into eight categories, where 0 = Early childhood education, 1 = Primary education, 2 = Lower secondary education, 3 = Upper secondary education, 4 = Post-secondary non-tertiary education, 5 = Short-cycle tertiary education, 6 = Bachelor's or equivalent level, 7 = Master's or equivalent level and 8 = Doctoral or equivalent level. The education level was skewed towards higher education levels; therefore, we only will compare those who have accomplished tertiary education (5 or higher on the ISCED scale) to those who have not. In the pooled sample, 250 respondents had a missing answer on the education questions. These respondents were recoded to have the mean education level of the relevant gender and origin country group, and we included a dummy variable for respondents with missing values.

Gender was treated as a grouping variable to compare means and coefficients between men and women. Different *origin countries* were treated as dummy variables (Bulgaria, Spain, Turkey) with Poland as a reference group.

Control variables. To control for changes over time, *months since migration* were calculated from respondents' month and year of arrival in the Netherlands. A*ge at first wave* was controlled since, for example, use of the host country language is greater for younger people which consequently might improve their labour market integration (Chiswick and Miller 2001). *Permanent settlement intentions* were controlled since migrants who intended to stay permanently were found to work fewer hours in an earlier study using the NIS-2NL data (Wachter and Fleischmann 2016). We compared those who expect to *stay* permanently in the Netherlands to those who expect to return to the country of origin, migrate to another country or are undecided. Lastly, being a *family migrant* was controlled since family migrants, who are disproportionally found among women, have lower labour market integration rates and being a family migrant is disproportionally more harmful to women's occupational status (Banerjee and Phan 2015). The respondents could choose from up to seven main reasons for their migration, and some migrants indicated both family and other reasons. The respondent was understood to be a family migrant if the primary reason for migration was mentioned to be at least one of the following: 'married someone already living in the Netherlands', 'joined other family members already living in the Netherlands' or 'moved together with family members'. Unless mentioned otherwise, all measures had no more than 2.8% missing values. In the analysis, cases with remaining missing answers in binary variables that were not imputed during data handling were dropped by list-wise deletion, while continuous variables with missing values were made endogenous, and therefore assumed to be multivariate normal distributed.

Analytical strategy

All main variables were measured in the same way in the three waves, which allows us to use them as time-varying variables. Only age at the first wave, ethnicity, level of education acquired in the origin country and settlement intentions at the first wave will be treated as time-invariant and included from the first wave responses only. Gender is used as grouping variable. All main analyses are conducted using Mplus (version 7; Muthén and Muthén 2007). As we analyse panel data, the observations are not independent but repeated within individuals at different time points. Therefore, we use multilevel growth models that take this nested data structure into account (Hox, Moerbeek, and van de Schoot 2010). The time-invariant variables are at the *individual level* (level 2) and the variables that are measured at different time points are at the *occasion level* (level 1).

Multilevel modelling also allows for incomplete observations, which means that respondents who have not participated in all three waves can still be included in the analysis. We use the Maximum Likelihood estimator to analyse the data. The option of random slopes will be utilised to test whether effects at the occasion level vary between individual respondents, and cross-level interactions will be included to test whether within-group variations can be predicted with individual-level characteristics. We will model both dependent variables separately and all model modifications will be done stepwise. A model is considered to have improved fit if both the deviance of the model and the Akaike Information Criteria (AIC) decrease compared to the previous nested model (Hox, Moerbeek, and van de Schoot 2010). Fit statistics of all tested models for both dependent variables are provided in Appendix 3 (Tables A4 and A5).

Results

Preliminary analysis

Descriptive statistics for all dependent, independent and control variables are presented in Table 1. The mean for *working hours* increased across waves, and analysis of variance

Table 1. Means and standard deviations for all variables in the analysis.

	Range	Wave 1 ($N = 3797$)	Wave 2 ($N = 1752$)	Wave 3 ($N = 1029$)
Hours worked[a]	0–60	37.04 (11.02)	37.19 (10.86)	38.09 (9.12)
Occupational status[a]	0–100	34.05 (22.15)	37.64 (23.80)	42.66 (23.98)
Individual level				
Tertiary education or higher	0–1	55.1%	–	–
Missing information about education	0–1	3.3%	–	–
Country or origin				
Bulgaria	0–1	18.0%	–	–
Poland	0–1	34.7%	–	–
Spain	0–1	30.3%	–	–
Turkey	0–1	16.9%	–	–
Female	0–1	55.7%	–	–
Age at wave 1	14–67	30.24 (8.17)	–	–
Permanent settlement intentions	0–1	34.0%	–	–
Family migrant	0–1	26.0%	–	–
Occasion level				
Months since migration	0–94	16.16 (12.52)	30.69 (12.24)	48.17 (12.37)
Labour market activity				
Working	0–1	52.1%	61.9%	69.3%
Studying	0–1	19.9%	14.2%	7.4%
Inactive	0–1	27.7%	23.9%	22.9%
Partner	0–1	51.6%	63.2%	15.2%
Children	0–1	17.5%	23.9%	35.6%
Missing information about household	0–1	13.2%	12.3%	0%
Traditional gender role values	1–5	1.94 (.88)	1.87 (.83)	1.84 (.77)

Note: *N* of group valid cases without missing answers.
[a]Only for respondents who are working.

revealed these changes to be statistically significant ($F(2) = 58.058$, $p < .001$). A Tukey post-hoc test showed that the mean significantly increased between all waves ($p < .001$). The respondents' mean for *occupational status* also increased across waves, and analysis of variance revealed that these changes were statistically significant ($F(2) = 27.299$, $p < .001$). A Tukey post-hoc test showed that the mean significantly increased between all waves (between *w1* and *w3* $p < .001$, between *w1* and *w2* $p < .05$, between w2 and *w3* $p < .01$). These results already underline the importance of studying recently arrived migrants' labour market integration with a longitudinal method, since significant changes occur in both measures between every wave.

Multilevel model

The baseline models for both dependent variables included only the time variable (*months since migration*) and the origin country dummies. For the baseline model of working hours, the dummy variables for main activity were also included.[10] Intra-class correlation coefficients indicated that among men, 5.2% of the variance in working hours and 73.5% of the variance in occupational status occur between individuals, and among women, 5.0% of the variance in working hours and 77.9% of the variance in occupational status occur between individuals. The remaining variation occurs within individuals over time.

In line with the descriptive statistics discussed above, time had a positive effect on the working hours of both men ($b = .043$, $SE = .010$, $p < .001$) and women ($b = .032$, $SE = .008$, $p < .001$). However, time did not have a significant impact on occupational status. This suggests that the descriptive trend of increasing occupational status over time is an artefact of origin group differences, which partly overlap with length of stay. When compared to the Polish reference group, men from Bulgaria worked fewer hours ($b = -2.481$, SE = .587, $p < .001$) but had higher occupational status ($b = 14.132$, SE = 2.442, $p < .001$). Men from Spain did not differ in working hours but had higher occupational status ($b = 31.875$, SE = 1.515, $p < .001$). Men from Turkey also worked fewer hours ($b = -1.633$, SE = .522, $p < .01$) and had higher occupational status ($b = 11.742$, SE = 1.837, $p < .001$). Among women, compared to the Polish reference group, those from Bulgaria worked fewer hours ($b = -2.281$, SE = .410, $p < .001$) but had higher occupational status ($b = 11.438$, SE = 2.194, $p < .001$). Women from Spain did not differ in working hours but had higher occupational statuses ($b = 30.658$, SE = 1.482, $p < .001$). Women from Turkey worked fewer hours ($b = -1.286$, SE = .458, $p < .01$) and had higher occupational status ($b = 23.155$, SE = 3.413, $p < .001$).

After estimating the baseline models, we added all predicting and control variables to test our hypotheses about main effects, and subsequently removed controls that were found to be insignificant.[11] We had hypothesised that *higher educated migrants work more hours* (H1a) and *have higher occupational status* (H7a). For men, only the latter hypothesis was accepted, since higher educated men had higher occupational status ($b = 12.984$, SE = 1.285, $p < .001$) but they did not work more hours than lower educated men. For women, both hypotheses were confirmed, since higher educated women worked more hours ($b = 1.121$, SE = .308, $p < .001$) and had higher occupational status ($b = 9.194$, SE = 1.414, $p < .001$).

Our hypotheses about the direct effect of having a partner on both dependent variables (H3a and H9) were both rejected as the pertaining regression coefficients were all insignificant, also when added without controlling for the presence of children. We, however, confirmed that *women with children work fewer hours* (H4a, $b = -1.068$, SE = .339, $p < .01$) and that *men with children work more hours* (H6a, $b = 1.258$, SE = .466, $p < .01$). Having children was found to be insignificant for the occupational status of both men and women, and therefore both hypotheses about this relationship (H10a and H11a) were rejected. We did not confirm that *women with traditional gender role values work fewer hours* (H5), but found that men with more traditional gender role values work fewer hours ($b = -.502$, SE = .198, $p < .05$). Additionally, we found traditional gender role values to result in lower occupational status for both women ($b = -2.836$, SE = .819, $p < .01$) and men ($b = -2.287$, SE = .660, $p < .01$).

Subsequently, we implemented Wald Tests of Parameter Constraints to test whether the regression coefficients and means of the outcome variables differ significantly between men and women. The effects which were not significantly different were constrained to be equal to increase parsimony. For *working hours*, significant differences for men and women were found in the effect of main activity ('in education' $Wald(1) = 17.768$, $p < .001$, 'inactive' $Wald(1) = 38.905$, $p < .001$) and the effect of having children ($Wald(1) = 16.276$ $p < .001$). Men and women had also significantly different means for working hours ($Wald(1) = 5.347$, $p < .05$) confirming our hypothesis (H2). For *occupational status*, significant differences for men and women were found in the effect of having higher education ($Wald(1) = 3.932$, $p < .05$) and being a family migrant ($Wald(1) = 11.578$, $p < .001$). Men and women did not have significantly different mean occupational statuses. Consequently, we rejected our hypothesis according to which *women have lower occupational status than men* (H7b) but accepted that *the effect of education is smaller for women than for men* (H8b), since the regression coefficient was smaller for women.

In the next step, we allowed the hypothesised occasion-level effects to vary within groups (random slopes) and tested whether the hypothesised individual-level characteristics could explain this variation (cross-level interactions). Some authors have recommended not to test cross-level interactions when no significant variation for the slope is found (Hox, Moerbeek, and van de Schoot 2010). However, others argue that in case cross-level interactions are hypothesised, these effects also should be tested despite the insignificance of the slope (LaHuis and Ferguson 2009). Therefore, even if we did not find significant random slopes, we tested the hypothesised cross-level interactions. Every time a new random slope was added also a covariance between the slope and the intercept, and between slopes were included (Hox, Moerbeek, and van de Schoot 2010).

A test for the random slope of time showed that men increased all in a similar manner in their working hours and occupational status. Also, women increased their working hours all in a similar manner but had significant variation in their occupational mobility over time ($b = .107$, $SE = .040$, $p < .01$). Since we hypothesised that *highly educated migrants have a smaller increase in their working hours* (H1b) but a *steeper increase in their occupational status* (H8a), we added educational level as a predictor for the random slope of time. Within the group of men, the cross-level interaction was found to be insignificant, but among women, opposite to what we hypothesised, highly educated women had a steeper increase in their working hours over time ($b = .026$, $SE = .013$, $p < .05$). Thus, hypothesis 1b was rejected for both men and women. Within the group of men, opposite to what was expected, the higher educated had a smaller increase in their occupational status ($b = -.118$, $SE = .053$, $p < .05$). For women, the cross-level interaction was insignificant. Thus, also hypothesis 8a was rejected for both men and women. In addition, we tested whether men and women from different origin countries change their working hours and occupational status at different rates. Compared to the Polish reference group and among the group of men, those from Spain increased their working hours marginally slower over time ($b = -.043$, $SE = .022$, $p < .1$). Among the group of women, those from Spain had a steeper increase in their occupational status over time ($b = .114$, $SE = .062$, $p < .05$). All other origin country differences were found insignificant.

Neither men nor women showed significant within-group variation in how having a partner affects their working hours or occupational status. However, we had hypothesised that *having a partner has a negative impact on women's working hours and a positive for men's especially if one has traditional gender role values* (H3b) and *less so if one is highly educated* (H3c). Therefore, interaction terms between having a partner and traditional gender role values and education were included. For men, the interaction with traditional gender role values was found insignificant. For women, opposite to what was expected, the interaction was significant and positive ($b = .562$, $SE = .283$, $p < .05$). Thus, hypothesis 3b was rejected for both men and women. Next, education was tested as a predictor for the random slope. This cross-level interaction turned out to be insignificant for both men and women, and therefore also hypothesis 3c was rejected. Additionally, we found that the effect of having a partner did not vary among men and women from different origin countries.

Both men and women showed significant within-group variation in how having children affected their working hours (men $b = 22.755$, $SE = 10.845$, $p < .05$, women $b =$

20.876, SE = 8.812, $p < .05$). Regarding occupational status, no significant variation was found. All our hypotheses that traditional gender role values would intensify the negative effects of motherhood and positive effects of fatherhood for one's working hours and occupational status (H4b, H6b and H10c) were rejected, since the pertaining interaction terms were found insignificant. Similarly, all hypotheses that expected the opposite if one had higher education (H4c, H6c, H10b and H11b) were rejected since the cross-level interactions with education were found insignificant. In addition, we tested whether the effect of children on both dependent variables would vary depending on the origin country. Compared to the Polish reference group and among the group of men, those from Turkey increased their working hours less if they had children ($b = -2.632$, SE = 1.229, $p < .05$). All other origin country differences were found insignificant.

Lastly, we also found some significant effects of the control variables. Migrants worked fewer hours if they intended to stay in the Netherlands permanently ($b = -.654$, SE = .241, $p < .01$) or had migrated as family migrants ($b = -1.890$, SE = .307, $p < .001$), but age did not play a significant role. Men who migrated as family migrants also had significantly lower occupational status ($b = -7.511$, SE = 1.797, $p < .001$). All other control variables for occupational status were found insignificant.

Table 2 shows the final model for both dependent variables, which are the most parsimonious models as they include equality constraints for means and coefficients that are not statistically different between men and women and only those random slopes that were found to be significant or had significant cross-level interactions. We can observe that men and women differ in their working hours in the first wave, but many of the predictors, including length of stay, work in the same way for both. Changes over time in working hours and household effects have some variation within the groups of men and women. Men and women have very similar trajectories in their occupational statuses and mainly differ in the returns to education, which are higher for men than for women. Occupational mobility over time shows some within-gender variation. The final model for working hours fits the data significantly better than the baseline model when comparing the change in deviance, but when comparing the AIC values, the model fit becomes worse ($\Delta\chi^2$ (Δdf = 45) = 393,500.790, ΔAIC = −13.262). However, as the AIC measure has a 'penalty function' for the increase in estimated parameters (Hox, Moerbeek, and van de Schoot 2010, 50), we argue that our final model still better explains the data than the baseline model. In our final model for occupational status, both fit indices are worse than in the baseline model ($\Delta\chi^2$ (Δdf = 22) = −14,641.766, ΔAIC = −14,713.766). This means that we were not able to specify a model that would explain recently arrived migrants' occupational status better than the first, empty baseline model did.

Additional analyses

As we use panel data in our analysis, we must consider the possibility that panel attrition is not random but selective. In other words, migrants with certain characteristics might have either moved out of the Netherlands or did no longer participate in the survey. This would mean that our results are not fully representative. We performed logistic regressions to estimate selective panel dropout for the sample that was analysed in our main analysis ($N = 6578$). We first tested whether panel dropout could be predicted by main activity, and found that those who were not working were less likely to participate in the following

Table 2. Final multi-group multilevel models of hours worked and occupational status for male and female immigrants including random slopes.

	Working hours per week		Occupational status	
	Men	Women	Men	Women
Intercept	39.376 (0.830)***	36.393 (0.757)***	26.117 (2.668)[a,***]	
Individual level				
Tertiary education	0.356 (0.396)[a]		17.530 (1.985)***	10.493 (1.945)***
Country or origin (ref = Poland):				
Bulgaria	−1.661 (0.565)[a,**]		13.361 (3.163)[a,***]	
Spain	−0.600 (0.492)[a]		22.408 (1.899)[a,***]	
Turkey	0.013 (0.574)[a]		16.512 (2.835)[a,***]	
Permanent settlement intentions	−0.654 (0.241)[a,**]		−0.719 (0.879)[a]	
Age	−0.014 (0.015)[a]		−0.061 (0.059)[a]	
Family migrant	−1.890 (0.307)[a,***]		−7.511 (1.797)***	−0.229 (1.501)
Occasion level				
Months since migration … [b]	[c]	[c]	[c]	[c]
Tertiary education[b]	0.005 (0.017)	0.030 (0.014)*	−0.130 (0.056)*	−0.064 (0.058)
Bulgaria[b]	−0.008 (0.024)	−0.023 (0.019)	−0.027 (0.090)	−0.001 (0.084)
Spain[b]	−0.031 (0.021)	−0.017 (0.019)	−0.015 (0.058)	0.144 (0.062)*
Turkey[b]	−0.031 (0.024)	−0.041 (0.023)[†]	−0.047 (0.080)	0.049 (0.115)
Intercept	0.054 (0.015)***	0.030 (0.012)*	0.125 (0.040)**	0.039 (0.042)
Main activity (ref = work)				
In education	−38.005 (0.523)***	−34.730 (0.405)***	*n.a*	
Inactive	−37.257 (0.492)***	−33.427 (0.319)***	*n.a*	
Partner (ref = no partner)	0.196 (0.735)[c]	−0.998 (0.601)[c,†]	−0.536 (0.783)[a]	
Children (ref = no children)	[c]	[c]	−0.521 (1.015)[a]	
… [b]				
Bulgaria[b]	0.842 (1.441)	0.600 (0.860)	*n.a*	
Spain[b]	0.399 (1.325)	−1.387 (0.865)	*n.a*	
Turkey[b]	−2.632 (1.229)*	0.177 (0.932)	*n.a*	
Intercept	1.953 (0.807)*	−1.048 (0.481)*	*n.a*	
Household answer missing	0.401 (0.337)[a]	5.219 (1.389)[a,***]	5.219 (1.389)***	
Traditional gender role values	−0.440 (0.182)[a,*]	−2.572 (0.513)[a,***]	−2.572 (0.513)***	
Partner × gender role values	−0.021 (0.310)	0.545 (0.286)[†]	*n.a*	
Random part				
Residual variances				
Occasion level *dv*	44.705 (2.508)***	36.633 (1.573)***	147.658 (15.417)***	89.895 (13.042)***
Individual level *dv*	22.107 (7.365)**	11.119 (4.120)**	184.831 (45.670)***	335.099 (50.658)***
Slope (*time*)	0.003 (0.006)	0.001 (0.004)	0.014 (0.030)	0.100 (0.040)*
Slope (*partner*)	7.488 (9.464)	7.772 (4.977)	*n.a*	
Slope (*children*)	23.577 (10.811)*	20.129 (8.858)*	*n.a*	
Covariances				
Individual level *dv* WITH				
Slope (*time*)	−0.125 (0.213)	0.014 (0.123)	−1.046 (1.118)	−3.660 (1.346)**
Slope (*partner*)	0.666 (7.387)	0.456 (3.834)	*n.a*	
Slope (*children*)	−15.147 (9.547)	−1.654 (4.978)	*n.a*	
Slope (*time*) WITH				
Slope (*partner*)	−0.098 (0.172)	0.073 (0.080)	*n.a*	
Slope (*children*)	0.037 (0.214)	−0.112 (0.100)	*n.a*	
Slope (*partner*) WITH				
Slope*(children)*	7.283 (6.550)	−11.978 (3.027)***	*n.a*	

Note: Unstandardized coefficients reported. Two-tailed significance tests *** $p < .001$, ** $p < .01$, * $p < .05$, [†] $p < .1$. *n.a* not applicable.
[a]Constrained to be equal across groups.
[b]Cross-level interaction.
[c]Random slope within groups.

waves. However, main activity only explained 0.4% of the dropout between the first and second wave and 2.8% between the second and third wave. When the dropout was predicted with all other variables used in the analysis, we found that those who had stayed in the Netherlands less time, did not have a partner or children, Polish, males and those who did not intend to stay permanently were less likely to participate a second time. Those who had a lower occupational status, had stayed in the Netherlands less time, did not have a partner or children, had a lower than tertiary education, Polish, older migrants and those who migrated for family reasons were less likely to participate the third time. With this model, we could explain 12.6% of the panel dropout between the first and second wave, and 74.7% between the second and third wave. These results are displayed in Appendix 4 (Tables A6 and A7). In sum, it seems that migrants with certain characteristic were more likely to drop out from the panel. Therefore, the findings about these relations should be interpreted with some caution. However, there is no reason to assume that the prevailing gender differences in labour market activity and strong gendered effect of children would be significantly biased by the selective panel attrition.

Furthermore, to test whether our results are robust to differences in length of stay and not dependent on the definition of *recently arrived*, we estimated the final model again for both dependent variables with a sample consisting of those who had stayed in the Netherlands for no longer than one year. We had two reasons for this. First, Amuedo-Dorantes and De la Rica (2007) found that among recently arrived migrants in Spain, those who have stayed in the country for 4–5 years were already significantly better integrated than those who had come within a year. Thus, some respondents in our main analysis might have been already so well integrated that the changes in their labour market integration during the survey would not be parallel to those who had stayed for a shorter time. Second, it is possible that migrants who arrived in a different period face different economic and political conditions, and consequently, their labour market trajectories might not be fully comparable (Busk et al. 2016). After the same data-handling process as for the sample in the main analysis, the sample sizes for the additional analysis were 1842 for the first wave, 826 for the second wave and 481 for the third wave. This sample is relatively small especially to perform multilevel multi-group analysis, which means that the results from our additional analysis should be interpreted with some caution. The results from both models are displayed in Appendix 4 (Table A8). The results from the additional analyses are in line with our main analyses, which indicate that our results are not dependent on how the boundary for recently arrived was defined. All the effects were in the same direction, only the level of significance was found lower for some of the effects in the alternative sample compared to the main analysis, but this can be probably accounted for the difference in the sample sizes.

Conclusion and discussion

This study has been one of the first to examine gender gaps in the labour market integration of recently arrived migrants, and moreover, how they develop during the first years after migration. Using multilevel multi-group modelling, we studied changes within persons over time, as well as differences between individuals over time and the causal relationships over time (Hox, Moerbeek, and van de Schoot 2010). This approach allowed us to study differences *between* men and women but also differences *within* the groups of men and women. Our findings confirmed that the labour market integration of recently arrived migrants, especially when it comes to labour market activity, show patterns that are disadvantageous to women, and thus similar to those found both among more established migrants as well as among natives in the Netherlands (e.g. Khoudja and Fleischmann 2015). Furthermore, as men and women change in their labour market activity at the same rate, women are not able to catch up their initial disadvantage over time. On the other hand, recently arrived male and female migrants are very similar when it comes to their occupational status, but they differ in the how they benefit from the education they received in their origin countries: in line with previous findings, returns to education were found to be substantially higher for men than women (Amuedo-Dorantes and De la Rica 2007).

Our study also contributed to the literature about the role of household composition by separating the effects of having a partner from those of having children, a consideration raised for example by Bevelander and Groeneveld (2012). Moreover, we studied men and women simultaneously, and took their ethnicity into account, whereas large streams of literature focus only on one of these. We did not find evidence that having a partner would positively influence the labour market integration of recently arrived migrants, a phenomenon that has often been found in cross-sectional studies among more established migrants and especially in relation to occupational mobility (Brekke 2013). Without actually controlling for partner characteristics, we assumed that the partners of recently arrived migrants would also be recently arrived, and consequently, to have limited abilities to help their partner's labour market integration via their own social networks. Future research should aim to better explore the relevant partner characteristics for recently arrived migrants' labour market integration. Contrary to our expectations, our results moreover showed that men who hold more traditional gender role values are less active on the labour market, whereas women who have a partner *and* traditional gender role values are actually more active on the labour market. This might be explained by the family investment hypothesis (Baker and Benjamin 1997), given the early phase of the integration process that we study. Future research that follows recently arrived migrants over a longer period is needed to study whether this initial gender gap reverses over time among the most traditional migrants.

The motherhood penalty and fatherhood premium for recently arrived migrants were not found regarding occupational status (Budig and Hodges 2010; Fleischmann and Höhne 2013), but with regard to labour market activity (Kaufman and Uhlenberg 2000). Contrary to what was expected, the strength of the motherhood penalty for recently arrived migrants is not affected by their level of education (Bevelander and Groeneveld 2012; England et al. 2016). This might be explained by the finding that the motherhood penalty is especially high for migrants with small social networks (Adsera and Chiswick 2007; Banerjee and Phan 2015), which is probably the case for many recently arrived migrants who have had relatively little time to establish networks in the Netherlands. Also, we did not find support for the claims that the motherhood penalty would be moderated by gender role values (Hakim 2000; Khoudja and Fleischmann 2015). This suggests that recently arrived migrants are not always able to put their preferences into practice when it comes to fitting family and work together.

When it comes to the differences between recent migrants from different origin countries, we found that men and women coming from the same origin country have very similar labour market integration trajectories. This is an important finding especially with regard to the established Turkish community in the Netherlands. The Turkish migrants are often pinpointed to have gendered labour market behaviour, and Turkish women in particular are found to have considerable difficulties in labour market integration (Khoudja and Fleischmann 2015). The new Turkish migrants seem to be better integrated and show fewer gender differences. Furthermore, migrants from Poland and Bulgaria are often studied jointly as Eastern European migrants. Our results show that there are clear differences between these groups, and treating them as one homogeneous group conceals important differences between them. Polish migrants are very active on the labour market, but at the same time, they have the lowest occupational statuses. By comparison, Bulgarians work relatively fewer hours, but at the same time, they hold quite high occupational statuses. Spanish migrants seem to be very active on the labour market, and also hold the highest occupational statuses. Especially among women, Spanish migrants experience relatively steep occupational mobility over time.

This research did not come without limitations. We could not control for selective panel attrition, which means that the respondents in later waves were not fully representative of the sample gathered at the first wave. We also were not able to separate between ageing and period effects. It could be that the changes over time were caused by migrants facing different economic and political restrictions at different times, and not so much by developments in their integration processes (e.g. Busk et al. 2016). The poor model fit for occupational status might be due to the fact that our sample had a relatively low share of respondents who were working and reported their occupation. Our sample was also rather homogeneous in terms of age and level of education, and due to the relatively small share of respondents with children, we could not separately study the effects of age and number of children (e.g. Budig and Hodges 2010). Future research should also focus on exploring labour market integration patterns of recently arrived migrants in other countries and test whether the current integration patterns we found in the Netherlands are more generalisable. It would be important to study countries where the motherhood penalty among natives is not as common as in in the Netherlands.

In sum, we contributed not only to the study of migrant labour market integration but also to the literature on gender inequality and the interlinked effects of gender, ethnicity, education and family composition. Especially, our longitudinal design is crucial in order to be able to gain more insights into the causalities in how migrants start their integration in a new country and how these patterns work differently for men and women. Our findings show that gender inequality is visible already shortly after arrival, but also that the measure used for integration makes a great difference in how these gender differences are visible. Integration policies should better take into account that men and women face different challenges for their labour market integration already upon arrival.

Notes

1. For exception, see Raijman and Semyonov (1997), Amuedo-Dorantes and De la Rica (2007) and Clark and Drinkwater (2008).
2. Data about the origin country and length of stay of the partner were available for this study, and initially we included this information in our analyses; however, we had to remove it again due to statistical power issues. As about 80% of the respondents with a partner reported the partner's origin country to be other than the Netherlands, our assumption that most partners are migrants themselves and therefore bound to have a limited social network in the Netherlands seems plausible.
3. Designed by Lubbers, Gijsberts, Fleischmann and Maliepaard (2015), carried out by Veldkamp Marktonderzoek BV and funded by NWO.

4. According to the municipal registry, between wave 1 and wave 2 (and between wave 2 and 3) 16.3% (9.8%) of the Bulgarians, 12.5% (7.5%) of the Polish, 24% (10.4%) of the Spanish and 10.4% (4.1%) of the Turkish had moved out of the Netherlands. Participants who moved within the Netherlands were contacted at their new address.
5. Five-year threshold adapted from Raijman and Semyonov (1997).
6. See Appendix 1, Table A1 for the share of men and women from each origin country per wave.
7. Twenty-three respondents in the first wave reported to have worked over 60 h per week, fifteen in wave 2 and four in wave 3. The largest number of weekly work hours reported was 93.
8. All respondents who had answered that their main activity is other than 'working' or 'in maternity leave' were recoded to have working hours equal to 0. This was necessary since the item about working hours asked about the current *or previous* job.
9. Previous studies about female labour market participation also suggested to take the age and number of children into account (Budig and Hodges 2010) but due to the low number of respondents with children in our sample (17.5–35.6%), there was insufficient statistical power for making these distinctions.
10. Activities 'in education' and 'inactive' were not possible to be included in the model to explain non-existing occupational status. They were tested both as predictors and covariates.
11. This concerns the control variable for cases with missing information about the level of education.

Disclosure statement

No potential conflict of interest was reported by the authors.

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Appendices

Appendix 1. Descriptive statistics

Table A1. Distribution of men and women from each country at each wave.

	Wave 1	Wave 2	Wave 3	Total
Bulgaria				
Males	417	200	125	742
Females	267	105	51	423
Poland				
Males	525	216	119	860
Females	794	384	240	1418
Spain				
Males	555	247	128	930
Females	597	304	170	1071
Turkey				
Males	335	157	104	596
Females	307	139	92	538
	3797	1752	1029	6578

Table A2. Share of missing answers in main analysis at each wave.

Working hours[a]	0.7%	0.9%	0.8%
Occupational status[a]	41.2%	57.9%	20.3%
Individual level			
Country of origin	0%	0%	05
Permanent settlement intentions	1.1%	1.2%	1.5%
Age at wave 1	0.4%	0%	0%
Family migrant	2.0%	2.2%	2.8%
Occasion level			
Months since migration	0%	0%	0%
Labour market activity	0.4%	0%	0.4%
Traditional gender role values	2.2	1.5%	1.7%

[a]Only for respondents who are working.

Appendix 2. Wald test of gender equality of effects in the main analysis

Table A3. Wald Test of gender equality of effects in the main analysis.

Main effect	Working hours	Occupation status
Mean	Wald(1) = 5.347 p = .0208	Wald(1) = 0.711, p = .3991
Individual level		
Tertiary education or higher	Wald(1) = 1.732, p = .1882	Wald(1) = 3.932, p = .0474
Country of origin		
Bulgaria	Wald(1) = 0.142, p = .7065	Wald(1) = 0.328, p = .5669
Spain	Wald(1) = 0.071, p = .7901	Wald(1) = 0.324, p = .5690
Turkey	Wald(1) = 0.144, p = .7048	Wald(1) = 1.337, p = .2475
Permanent settlement	Wald(1) = 1.553, p = .2126	Wald(1) = 0, p = 1
Age	Wald(1) = 1.606 p = .2051	Wald(1) = 0.036, p = .8501
Family migrant	Wald(1) = 0.448, p = .5032	Wald(1) = 11.578, p = .0007
Occasion level		
Months since migration	Wald(1) = 0.120 p = .7295	Wald(1) = 0.860, p = .3536
Main activity (ref = work)		
In education	Wald(1) = 17.768, p = .000	*n.a*
Inactive	Wald(1) = 38.905, p = .000	*n.a*
Living with partner	Wald(1) = 0.004, p = .9514	Wald(1) = 0.933, p = .3342
Living with children	Wald(1) = 16.276 p = .0001	Wald(1) = 1.698, p = .1925
Missing information on household	Wald(1) = 2.775, p = .0957	Wald(1) = 0.010, p = .9187
Traditional gender role values	Wald(1) = 2.302, p = .1292	Wald(1) = 0.255, p = .6138

Note: *n.a*: not applicable.

Appendix 3. Model fit statistics

Table A4. Goodness-of-fit statistics of all the tested models for dependent variable working hours.

Model	Model description	*N* of free parameters	Deviance	Δ Deviance	AIC
M1	Baseline model	18	451,858.5	–	45,221.845
M2	All predictors added	38	59,307.38	392,551.1	59,383.383
M3	Equality constrains	27	59,318.78	−11.398	59,372.782
M4	Random slope time	32	59,319.79	−1.012	59,383.793
M5	Cross-level interaction (education)	34	59,315.1	4.692	59,383.101
M6	Cross-level interaction (country of origin)	40	59,309.36	5.746	59,389.355
M7	Random slope partner	47	59,291.18	18.176	59,385.181
M8	Partner × gender role values	49	58,411.02	880.162	58,509.018
M9	Cross-level interaction (education)	51	58,410.97	880.214	58,512.965
M10	Cross-level interaction (country of origin)	55	58,404.7	6.318	58,514.700
M11	Random slope children	57	58,369.73	939.624	58,483.732
M12	Children × gender role values	59	58,369.49	0.24	58,487.493
M13	Cross-level interaction (education)	59	58,366.61	3.126	58,484.605
M14	Final model: cross-level interaction (country of origin)	63	58,357.71	12.026	58,483.706

Note: Model fit compared to the previous nested model.

Table A5. Goodness-of-fit statistics of all the tested models for dependent variable occupational status.

Model	Model description	*N* of free parameters	Deviance	Δ Deviance	AIC
M1	Baseline model	14	18,845.46	–	18,873.464
M2	All predictors added	34	33,529.89	−14,684.4	33,597.889
M3	Equality constrains	23	33,540.57	−10.686	33,586.574
M4	Random slope time	28	33,527.41	13.16	33,583.414
M5	Cross-level interaction (education)	30	33,522.46	4.952	33,582.461
M6	Cross-level interaction (country of origin)	36	33,515.23	7.232	33,587.230
M7	Random slope partner	43	33,505.45	9.776	33,591.454
M8	Cross-level interaction (country of origin)	49	33,500.92	4.53	33,598.925
M9	Random slope children	41	33,515.26	−0.03	33,597.260
M10	Children × gender role values	45	33,176.28	338.984	33,266.275
M11	Cross-level interaction (education)	45	33,515.1	0.156	33,605.104
M12	Cross-level interaction (country of origin)	49	33,510.5	4.758	33,608.502
M13	Final model (Model 6)	36	33,515.23	4.952	33,587.230

Note: Model fit compared to the previous nested model.

Appendix 4. Additional analysis

Table A6. Logistic regression coefficients and standard errors for panel dropout, predicted by main activity.

	Panel dropout between wave 1 and wave 2	Panel dropout between wave 2 and wave 3
Main activity (ref = working)		
Studying	0.214 (.081)**	1.123 (.126)***
Inactive	0.214 (.067)**	0.372 (.082)***
Fit statistics		
Nagelkerke R^2	.004	.028

Note: Two-tailed significance tests. *$p < .05$, **$p < .01$, ***$p < .001$, $^{\dagger}p < .1$.

Table A7. Logistic regression coefficients and standard errors for panel dropout, predicted by all dependent and predicting variables.

	Panel dropout between wave 1 and wave 2	Panel dropout between wave 2 and wave 3
Hours worked	−0.001 (.006)	0.006 (.009)
Occupational status	0.000 (.003)	−0.012 (.005)*
Months since migration	−0.014 (.003)***	−0.153 (.008)***
Partner	1.433 (.130)***	3.477 (.266)***
Children	0.447 (.151)**	−2.654 (.267)***
Tertiary education or higher	−0.076 (.136)	−1.043 (.216)***
Country of origin (ref = Poland)		
Bulgaria	−0.503 (.214)*	−1.354 (.303)***
Spain	−0.595 (.177)**	−1.354 (.268)***
Turkey	−0.750 (.217)**	−1.428 (.335)***
Female	−0.285 (.126)*	0.145 (.187)
Age at wave 1	−0.006 (.008)	0.025 (.012)*
Permanent settlement intentions	−0.217 (.117)†	−0.137 (.180)
Family migrant	0.002 (.153)	−1.244 (.243)***
Fit statistics		
Nagelkerke R^2	0.126	0.747

Note: Two-tailed significance tests. *$p < .05$, **$p < .01$, ***$p < .001$, $^{\dagger}p < .1$.

Table A8. Final multi-group multilevel models of working hours and occupational status for male and female migrants including random slopes, additional analysis among migrants who had migrated within a year before the first wave.

	Working hours per week		Occupational status	
	Men	Women	Men	Women
Intercept	38.842 (1.142)***	35.019 (1.055)***	27.993 (4.348)[a],***	
Individual level				
Tertiary education	0.189 (0.483)[a]		14.889 (2.641)***	5.288 (2.728)†
Country of origin (ref = Poland)				
Bulgaria	−1.870 (0.709)[a],**		15.584 (4.273)[a],***	
Spain	0.154 (0.600)[a]		22.408 (2.547)[a],***	
Turkey	0.484 (0.695)[a]		15.657 (3.690)[a],***	
Permanent settlement intentions	−0.753 (0.351)[a],*		0.975 (2.641)[a],***	
Age	−0.009 (0.022)[a]		−0.016 (0.102)[a]	
Family migrant	−2.268 (0.433)[a],***		−8.811 (2.533)**	−0.527 (2.347)
Occasion level				
Months since migration … [b]	[c]	[c]	[c]	[c]
Tertiary education[b]	0.036 (0.033)	0.059 (0.029)*	0.099 (0.114)	0.242 (0.131)†
Bulgaria[b]	−0.069 (0.052)	0.012 (0.040)	−0.235 (0.182)	−0.271 (0.174)
Spain[b]	−0.110 (0.041)**	−0.030 (0.033)	−0.125 (0.119)	0.099 (0.115)
Turkey[b]	−0.065 (0.046)	−0.041 (0.040)	−0.230 (0.146)	−0.008 (0.193)
Intercept	0.108 (0.040)**	0.021 (0.031)	0.153 (0.114)	
Main activity (ref = work)				
In education	−38.338 (0.663)***	−34.292 (0.535)***	*n.a*	
Inactive	−36.223 (0.664)***	−32.469 (0.462)***	*n.a*	
Partner (ref = no partner) … [b]	−0.112 (0.980)[c]	−0.480 (0.858)[c]	0.644 (1.354)[a]	
Children (ref = no children)	[c]	[c]	−0.356 (1.705)[a]	
Bulgaria[b]	1.365 (1.920)	3.499 (1.383)*	*n.a*	
Spain[b]	2.585 (1.745)	−1.013 (1.250)	*n.a*	
Turkey[b]	−1.595 (1.593)	−0.202 (1.317)	*n.a*	
Intercept	0.286 (1.227)	−0.855 (0.703)	*n.a*	
Household answer missing	0.844 (0.458)[a],†	8.783 (2.376)[a],***	8.783 (2.376)***	
Trad. gender role values	−0.320 (0.248)[a]	−3.096 (0.855)[a],***	−3.096 (0.855)***	
Partner*gender role values	−0.091 (0.429)	0.360 (0.421)	*n.a*	
Random part				
Residual variances				
Occasion level dv	35.862 (3.303)***	34.428 (2.539)***	115.290 (26.213)***	109.523 (26.132)***
Individual level *dv*	30.463 (8.060)***	6.832 (4.924)	192.864 (57.630)**	312.201 (71.017)***
Slope*(time)*	0.017 (0.013)	0.009 (0.010)	0.066 (0.075)	0.136 (0.084)
Slope*(partner)*	2.555 (14.920)	9.022 (7.762)	*n.a*	
Slope*(children)*	6.088 (12.064)	20.680 (13.535)	*n.a*	
Covariances				
Individual level *dv* WITH				
Slope*(time)*	−0.503 (0.313)	0.085 (0.207)	−1.531 (1.838)	−3.042 (2.192)
Slope*(partner)*	3.982 (10.588)	1.720 (5.629)	*n.a*	
Slope*(children)*	−11.705 (16.802)	−1.743 (6.703)	*n.a*	
Slope*(time)* WITH				
Slope*(partner)*	0.038 (0.283)	0.212 (0.167)	*n.a*	
Slope*(children)*	0.230 (0.387)	−0.294 (0.211)	*n.a*	
Slope*(partner)* WITH				
Slope*(children)*	−0.925 (13.340)	−12.635 (4.506)**	*n.a*	

Notes: Unstandardized coefficients reported. Two-tailed significance tests. *$p < .05$, **$p < .01$, ***$p < .001$, †$p < .1$.
n.a: not applicable.
[a]Constrained to be equal across groups.
[b]Cross-level interaction.
[c]Random slope within groups.